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To my family

Preface

Unbelievably, but I am really writing the preface of my doctoral thesis. It seemed fast at last, but I remember many situations where the finalisation of this project seemed far away. That I am now completing my dissertation is something that I owe to many people.

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This research has benefited a lot from comments and remarks from anonymous referees and editors of various scientific journal as well as from comments at EALE/SOLE world conference 2005 in San Francisco, EALE conference 2006 in Prague, Verein für Socialpolitik 2005 in Bonn and 2006 in Bayreuth, the Interdisciplinary Spatial Statistics Workshop in Paris (JISS) 2004, the IAB Nutzertagung 2005 as well as from comments at seminars at TU Darmstadt, Goethe-University Frankfurt and Leicester University - thank you!

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Mannheim,

February 2007 *Melanie Arntz*

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Introduction

In the OECD employment outlook 2005 (OECD, 2005), a comparison of internal migration rates[1] across countries confirms a well-known stylised fact in empirical economics[2]: internal migration rates are lower in European countries than in the US or Australia. Moreover, eastern and southern Europe have the lowest and western Europe, especially the UK, have the highest migration rates in Europe, while Germany ranks in between these extremes. This current level of internal migration in Germany results from rising mobility rates since the end 1990s after a period of stagnation in western Germany during the 1970s and 1980s (see Entorf, 1996; Faini, 1999; Haas, 2000). Net migration from eastern to western Germany peaked around re-unification and has been dropping markedly until the mid 1990s (Maretzke, 1998), a trend that has been associated with a rapid wage convergence in the early 1990s (Hunt, 2000). Since the mid 1990s, however, the pace of economic recovery of eastern Germany has been slowing down and east-west migration has again started to rise (Werz, 2001; Heiland, 2004).

Understanding the determinants of these observed levels of interregional mobility is of major policy concern because the geographic mobility of labour may contribute to higher overall economic growth as well as a reduction of interregional employment disparities. Higher economic growth arises if interregional mobility reduces the regional mismatch between labour supply and labour demand and thus leads to higher employment levels. Such potential welfare gains from geographic mobility are likely to be substantial for Germany because the share of unemployment caused by a regional mismatch is estimated to be 20% (SVR, 1994)

[1] Throughout the thesis, internal migration, interregional mobility and geographic mobility are used synonymously and always refer to the relocation of residence across regional boundaries within a country. The exact distinction of intraregional and interregional movements is presented in the subsequent chapters.

[2] See also Eichengreen (1991), De Grauwe and Vanhaverbeke (1993), Faini (1999), and Braunerhjelm et al. (2000).

or even 45% (Entorf et al., 1990). Moreover, geographic mobility may also foster productivity and thus economic growth by improving the average matching quality. This is because the quality of a job match should increase with the size of the labour market, i.e. the number of accessible workplaces. Since geographic mobility extends labour markets beyond the local labour market area, it may thus also raise the average matching quality (Kim, 1989).

In addition, internal migration may be an important means of equilibrating regional employment disparities. According to Blanchard and Katz (1992), an adverse region-specific labour demand shock in the US initially raises unemployment and reduces nominal wages and participation rates. While lower wages stimulate labour demand and thus offset some of the initial shock, the main adjustment occurs via net out-migration. This adjustment process responds very quickly so that unemployment and participation rates already return to the pre-shock level after 5 to 7 years. Similarly, internal migration has also been found to be a well-functioning means of adjustment in Australia (Debelle and Vickery, 1999) and in New Zealand (Choy et al., 2002). In contrast, the responsiveness of internal migration to employment shocks seems to be much less pronounced in Europe. Decressin and Fatàs (1995) as well as Nahuis and Parikh (2002) examine regional adjustment dynamics in Europe and find that internal migration responds much slower to a negative demand shock than in the US. Therefore, the main adjustment in depressed regions rather occurs via lower participation rates. By contrast, Möller (1995) suggests that participation reacts weakly to a regional shock in Germany and that migration is a mechanism for regional adjustment at least in the long run. A common finding in all these studies is that migration does not seem to respond to regional shocks in the short run in Europe. This result seems to be in line with Puhani (2001) who suggests lower elasticities of aggregate migration flows with respect to unemployment and wage differentials in Europe and Germany than in the US. As a consequence of such weak or slow adjustment processes, regional employment disparities are likely to be quite persistent in Europe. This is confirmed for many European countries (Martin, 1998). The range between the region with the highest and the lowest unemployment rate often exceeds ten percentage points. In particular, Italy and Germany are characterised by strong regional unemployment disparities which coincide with the south-north divide in the case of Italy and the east-west divide in the case of Germany (OECD, 2005).

A higher level of geographic mobility in Germany could thus result in a number of po-

tential welfare gains. First of all, the estimated extent of regional mismatch in Germany indicates some scope for increasing employment levels and reducing unemployment. Secondly, higher productivity levels in case of an increasing matching quality could additionally contribute to higher economic growth. And finally, higher mobility levels could accelerate adjustment processes after region-specific shocks and thus reduce regional employment disparities. Removing barriers to interregional mobility is thus an important policy issue in order to realise these potential welfare gains. However, the degree of heterogeneity in internal migration rates across different labour market segments suggests that identifying obstacles to the mobility of labour may necessitate a closer look at heterogeneous individuals. As an example, Figure 0.1 shows the shares of interregional job movers among all job movers by educational attainment and type of job change in western Germany. In particular, less-skilled individuals tend to be less mobile than their high-skilled counterparts. Moreover, direct job-to-job movers (DJC) are much more mobile than their previously unemployed counterparts who received unemployment compensation prior to the job move (JCU).

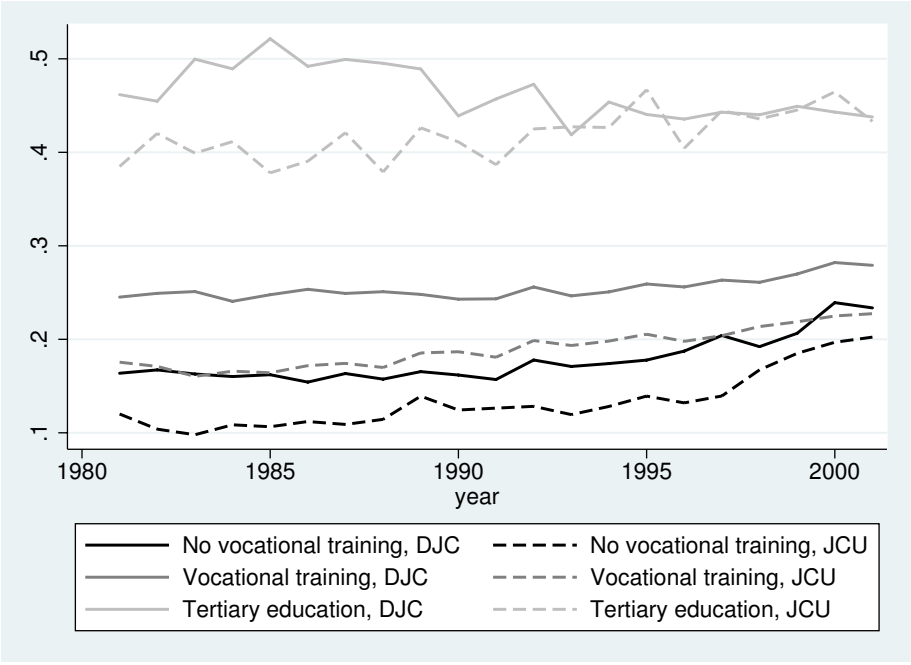


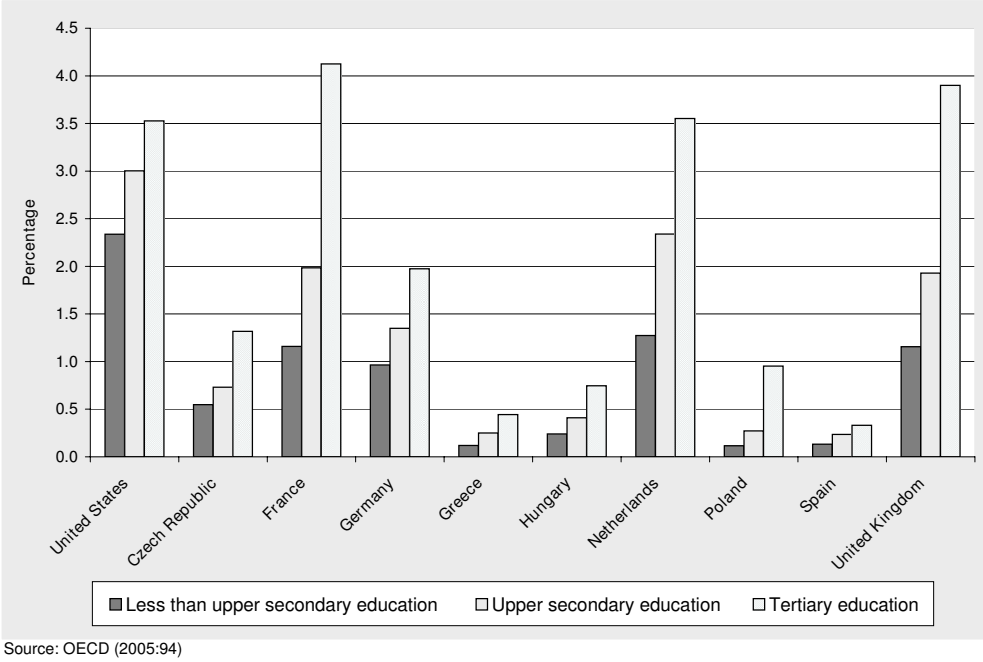
Figure 0.1: Share of interregional job moves by type of job change and educational attainment in western Germany, 1981-2001 (Source: Own calculation based on IAB-R01 (see chapter 5 for details))

These differences across groups suggest that migrants must not be treated as a homogeneous group, but rather have to be understood as individuals with a unique disposition to be mobile. This individualised perception of migrants has first been introduced by the human capital model of migration (Sjaastad, 1962; Shields and Shields, 1989). According to this framework, potential migrants compare the expected future returns in the destination area to the expected future returns when staying in the current place of residence and move if the former exceeds the latter. Since individuals may differ in terms of their time horizon, their preferences and their potential to increase future returns by migrating, individuals may be differently prone to migration. Intervening obstacles or barriers to mobility could thus be described as institutional factors that reduce the expected returns to migration or increase migration costs and thus reduce an individual's probability of migration. As an example, being unemployed is typically assumed to increase the probability of migration. In fact, several empirical studies for the US and the UK find higher migration propensities for unemployed compared to employed individuals (Herzog and Schlottmann, 1984; Pissarides and Wadsworth, 1989; Jackman and Savouri, 1992; Bailey, 1993). The unemployment compensation and welfare system, on the other hand, might increase the costs of migration by increasing the returns to staying unemployed in the place of residence and might thus generate an obstacle to mobility for unemployed benefit recipients (Hassler et al., 2005). In line with this notion, a Spanish study by Antolin and Bover (1997) indicates lower mobility levels among unemployed benefit recipients than among non-recipients. In addition to compositional effects, lower mobility levels for job movers after an unemployment period in Figure 0.1 may thus also reflect the disincentive effect of the unemployment compensation system because this group differs from direct job movers by the previous receipt of unemployment compensation.

Other institutional factors that may be considered to shape individual migration propensities and to affect different labour segments to a varying extent are, for example, active labour market policies, housing policies, and collective wage bargaining. Participating in active labour market programs, for example, might enable individuals to postpone migration. An extensive local supply of such programs may thus result in a regional locking-in effect (Frederiksson, 1999; Westerlund, 1998), especially among those for whom such programs are an attractive substitute for regular employment. Moreover, individuals may be quite differ-

ently affected by housing policies. For homeowners, transaction costs in housing markets and possible capital losses have been associated with increasing migration costs and thus lower mobility levels among homeowners (van Ommeren, 1996; Henley, 1998; Barcelo, 2003). For individuals at the fringe of the labour market, however, social housing may be more relevant as a barrier to mobility because access to social housing may not be easily moved from one region to the other. Empirical studies confirm such disincentive effects of social housing for several European countries (Gardner et al., 2001; Barcelo, 2003). In addition, interregional wage disparities are likely to be small if central wage bargaining restricts the scope for local wage agreements (OECD, 2004). Such wage rigidities limit the returns to migration with respect to potential wage gains and thus confine migration. The resulting disincentives to migration should, however, be less severe for high-skilled individuals who are more likely to earn an income outside the collective wage agreement.

Institutional factors may thus have a varying impact on different segments of the labour market. If the unemployment compensation and welfare system, for example, proves to be an obstacle to interregional mobility, this obstacle should be more pronounced for labour market segments which are highly dependent on transfer payments. Institutional factors may thus be partially responsible for the relatively low responsiveness of less-skilled individuals to regional shocks in Spain compared to their high-skilled counterparts (Mauro and Spilimbergo, 1999). Moreover, institutional cross-country differences may then provide one explanation why migration rates across skill groups differ across countries. Figure 0.2 shows internal migration rates by educational attainment for several OECD countries. Note that Figure 0.2 displays migration rates among the total population aged 15 to 64, while Figure 0.1 only referred to job movers. Since job movers are likely to be a relatively selective and geographically mobile group, this may explain the much higher share of interregional moves in Figure 0.1 as compared to Figure 0.2. Still, Figure 0.2 again confirms the lower level of geographic mobility among less-skilled individuals. Moreover, low-skilled individuals in Europe seem to be even less mobile compared to their high-skilled counterparts than in the US. Such differences may reflect institutional cross-country differences not only in the labour market, but also, for example, in housing policies.



Source: OECD (2005:94)
Note: Residential change refers to NUTS2 level for European countries and to the state level for the US.

Figure 0.2: Proportion of persons aged 15-64 who changed residence in 2003

Removing institutional barriers to mobility in order to promote geographic mobility, however, can never be an end in itself because a higher level of interregional mobility may also result in welfare losses that need to be taken into account. First of all, social costs from a higher level of geographic mobility may arise due to a weakening of family and friendship ties that aggravates the informal organisation of the care for the elderly as well as child care and may thus also be related to lower fertility levels. Secondly, rising rents in prospering regions that attract net inward migration may result in a displacement of socially weaker groups which may intensify social tensions (Heye and Odermatt, 2006). Another negative consequence of geographic mobility are the induced adjustment costs as the provision of public goods such as schools, hospitals, public transport etc. needs to be adjusted to a changed level and composition of local residents. As an example, the continued net loss of population in eastern Germany necessitates a downsizing of public spending on schools and public transport as well as a dismantling of vacant housing space (IWH, 2006). Moreover, the negative selection of those remaining in depressed regions in terms of qualification and age may undermine the future growth potential of a region. This is because the local availability of a large pool of qualified workers is vital for regional innovation and thus the endogenous growth potential of the region (Lucas, 1988; Romer, 1990). Therefore, even if migration goes

from high to low unemployment regions and from low to high wage regions, migration need not necessarily reduce regional disparities as postulated by the neo-classical model. Instead, the composition of net migration flows in terms of qualification and other productivity related characteristics may strongly influence whether migration contributes to regional convergence or rather fosters regional divergence.

The objective of this thesis thus is a twofold. On the one hand, the aim is to shed light on the determinants of mobility for heterogeneous labour market segments in order to identify the scope for increasing geographic mobility in Germany. In particular, the thesis focuses on the determinants of mobility of unemployed individuals because the willingness and ability of this labour market segment to seek employment elsewhere is of central concern if migration is to contribute to reducing regional disparities and increasing employment levels. A major part of this thesis thus looks at the responsiveness of unemployed individuals to regional labour market conditions and the incentives and disincentives to mobility caused by labour market institutions such as active and passive labour market policies. The aim is to identify possible barriers to mobility which may explain the observed low mobility level in Germany, especially among even less mobile low-skilled unemployed. In addition, another objective of this thesis is to examine the determinants of the skill composition of migration flows in Germany. By doing so, the thesis identifies the scope for shaping the selectivity of migration and thus shows ways how to support the convergent rather than the divergent character of geographic mobility. It thus provides insights on how to mitigate one possible drawback of a higher level of geographic mobility.

The theoretical foundation for these analyses mainly comes from two approaches to the microeconomics of migration. On the one hand, assuming individuals to be heterogeneous optimisers of their future expected gains is clearly informed by the human capital model of migration. On the other hand, search theory provides another important theoretical foundation for the analyses presented in the following chapters. In particular, individuals are not assumed to have perfect information on all possible job alternatives. Instead, individuals have to search for jobs by deciding on the optimal spatial search strategy (Herzog et al., 1993). Both of these theoretical frameworks echo through this thesis.

The thesis consists of five separate papers, partially written with co-authors. The micro data used for the subsequent empirical analysis of individual mobility behaviour are the

IAB employment subsamples regional files 1975-1997 and 1975-2001 (see 2.3 and 5.3 for details) as well as the Integrated Employment Biographies V.1 (see 3.4 for details). All these administrative data allow for a reconstruction of individual employment histories on a daily basis including information on the workplace location and thus geographic mobility. This information and the size of the data representing 1-2.2% of the population working in a socially insured job provide a data base that is well-suited for the analysis of individual mobility behaviour even for rather small sub-groups of the labour force. One problem that is common to all these micro data, however, is that workplace locations are coded at the level of microcensus regions while major labour market attributes that need to be combined with the micro data for the subsequent analyses are coded at the level of employment agency districts. Since these two sets of regions are geographically incompatible, i.e. one set of regions does not in general respect the boundaries of the other set and the two sets are not nested hierarchically, an appropriate interpolation technique is needed in order to transfer attributes from employment agency districts to counties and vice versa. **Paper (1)** presents a solution to this areal interpolation problem. In particular, paper (1) considers different cartographic interpolation methods such as simple area weighting and dasymetric weighting both of which are based on estimated intersection areas. The preparatory work in paper (1) thus builds up a data base for the following empirical papers that combines the micro data with interpolated region-specific attributes.

The next three papers focus on the determinants of mobility for unemployed individuals and examine the responsiveness of unemployed individuals to regional labour market conditions. Since labour market institutions[3] such as the unemployment compensation system are likely to affect mobility decisions of unemployed jobseekers, papers (2) to (4) also explore to what extent such institutions create barriers to mobility for this labour market segment.

Paper (2) examines mobility decisions of unemployed individuals in western Germany between 1983 and 1997. Based on a job search model with multiple regional labour markets, this paper investigates whether unemployed individuals in western Germany choose search strategies that favour migration away from depressed regions. Moreover, the paper also looks at the responsiveness to local labour market conditions among different labour market

[3]Since the empirical results in this thesis have all been obtained for a period prior to the latest Hartz reforms, labour market policies refer to the previous institutional setting.

segments such as low-skilled and skilled individuals and takes account of some passive and active labour market institutions that may affect individual mobility decisions. The empirical analysis uses a competing-risk hazard framework of exiting unemployment to jobs in a local or a distant labour market area. Estimation results are obtained from a stratified Cox proportional hazard model that allows for region-specific fixed effects. This approach reduces biases due to unobserved regional heterogeneity (e.g. unobserved regional amenities) that affects both the migratory behaviour of jobseekers and the regional conditions of interest. Estimates based on this stratified approach indicate that at least high-skilled men are responsive to local labour market conditions and have higher migration probabilities in regions with unfavourable re-employment opportunities. By contrast, women and low-skilled men tend to stay in regions despite relatively unfavourable job-finding prospects and are thus more dependent on local labour market conditions. The results also indicate that passive labour market measures seem to be an important determinant of the migratory behaviour of unemployed jobseekers. In particular, long entitlements to unemployment benefits strongly reduce mobility among all jobseekers. By contrast, an extensive local supply of work creation schemes only seems to have a weak regional locking-in effect on women. The interpretability of this latter finding is, however, limited since the micro data does not allow for distinguishing exits to the regular labour market from exits to subsidised employment in the context of those labour market programs.

In several respects, **paper (3)** supplements the previous paper. Based on the latest generation of German administrative micro data, regular employment can now be distinguished from subsidised employment in the context of a labour market program (e.g. work creation schemes). Compared to the previous paper, paper (3) thus extends the competing risk set to also include exits to subsidised employment. The objective then is to explore the main determinants of the length of unemployment and to disentangle the relevance of individual, regional and institutional factors for exiting unemployment to one of the three competing exit states. In particular, the paper provides evidence about the extent to which passive and active labour market policies as well as local economic conditions and job counselling activities affect the duration of unemployment and the resulting labour market state. Paper (3) thus differs from paper (2) not only with regard to the set of competing risks, but also explores the impact of a broader set of regional and institutional factors on the unemploy-

ment experiences of individuals not only in western, but also in eastern Germany between 2000 and 2004. Moreover, the paper draws specific attention to unemployment experiences of individuals with low pre-unemployment wages because this group is likely to be less mobile and to experience a prolonged unemployment period. Distinguishing between regular and subsidised employment reveals that subsidised employment often cushions unfavourable local labour market conditions for less mobile labour market segments such as married men and low-earning individuals whereas well-earning singles rather experience higher migration levels. In line with the previous paper, there is again significant evidence that long entitlements to unemployment benefits prolong the duration of unemployment and reduce migration, especially among well-earning unemployed for whom exhausting unemployment benefits entails some major reductions in transfer receipt. By contrast, other institutional factors such as active labour market programs and local job placement activities only seem to have a minor impact.

The previous two papers point towards the role of passive labour market measures in affecting labour market outcomes of unemployed jobseekers. Long entitlements to unemployment benefits (UB) have been found to prolong unemployment and reduce migration. Interpreting these findings as a causal relationship may be misleading, however, if the results are partially driven by unobserved factors of the working history that determine individual UB entitlements and also affect individual job-finding chances. **Paper (4)** thus re-examines the effect of unemployment compensation on the migratory behaviour of unemployed jobseekers by using variation in the length of entitlements to unemployment benefits from a natural experiment that is provided by a labour market reform in 1997. By comparing transitions to local and non-local employment in a pre-reform and post-reform period for a treatment group whose UB entitlements have been cut after 1997 to a control group whose UB entitlements have been left unaffected, it is possible to identify the causal effect of unemployment benefits on migration under reasonable assumptions. As a major contribution to the literature, paper (4) presents an approach how to analyse treatment effects on competing failure types in the case of partially missing information on the latent failure times. Partially missing information occurs because unobserved periods in an individual's employment record in the German administrative data result in incomplete knowledge concerning the duration until leaving unemployment to competing failure types such as self-employment or

leaving the labour force. In the context of dependent competing failure types, bounds for the treatment effect on the marginal survivor curves may be biased. The paper thus proposes a non-parametric bounds analysis of risk-specific cumulative incidence curves (CIC) to bound the treatment effect on the CIC over different definitions of the latent durations. Although this approach does not resolve the non-identifiability of competing risks, it provides a generally applicable and flexible descriptive tool for the observed distribution of competing failures in the case of dependent competing risks. For high-skilled individuals, for whom the threat of entitlement loss due to the 1997 reform is likely to be largest, the bounds for the cumulative incidence of migration are indicative for the mobility-reducing effect of extensive UB receipt. A relatively generous unemployment compensation in Germany may thus contribute to lower migration levels in Germany as compared with, for example, the US.

The analysis of the migratory behaviour of unemployed jobseekers provides some insights into the responsiveness of this labour market segment to regional disparities as well as the barriers stemming from certain labour market institutions. Papers (2) to (4) thus help in identifying the scope for policy makers to shape mobility levels and to realise some of the aforementioned welfare gains from a higher level of mobility. As has previously been discussed, identifying the scope for increasing geographic mobility is only one important aspect to look at that needs to be complemented by an assessment of its negative consequences and possible ways to cushion them. The last paper, **paper (5)**, thus wants to shed light on how to mitigate one particular downside of geographic mobility, namely the possibly divergent effect of internal migration on the regional system. In the German context, the net migration of high-skilled individuals from eastern to western Germany raises strong concerns that a brain drain from eastern to western Germany may undermine the future growth potential of eastern Germany and may thus reinforce regional east-west disparities. Paper (5) thus looks at the factors that determine the skill composition of internal job matching flows in Germany. For this purpose, the analysis examines destination choices of heterogeneous skill groups. Moreover, the paper investigates whether different destination choices of direct job-to-job changers and job changers after an unemployment period contribute to the observed spatial job matching pattern by skill group. Based on a sample of job moves between 1995 and 2001, estimates are derived from a nested logit model that takes account of unobserved interregional heterogeneity. The findings indicate that spatial job matching patterns by high-

skilled individuals are mainly driven by interregional income differentials, while interregional job matches by less-skilled individuals are mainly affected by interregional differentials in job-finding opportunities. Interregional amenity differentials only weakly contribute to spatial sorting processes in Germany. Such differences in destination choices by skill level seem to be partly modified by different spatial patterns of job-to-job matches and job matches after unemployment. A simulated economic convergence between eastern and western Germany demonstrates that higher wage levels are the most effective means of attracting human capital to eastern Germany, but that the net loss of population can only be reversed by lower unemployment rates. If maintaining the future viability of eastern Germany is a pronounced policy objective, the findings thus advocate policies that foster wage convergence without further increasing eastern unemployment levels. Paper (5) thus provides insights on how policy can promote integration and convergence between eastern and western Germany.

Finally, the last section of this thesis contains a number of concluding remarks and an outlook on future research needs.

Chapter 1

An Application of Cartographic Area Interpolation to German Administrative Data

joint with Ralf A. Wilke[1]

Abstract

In many situations the applied researcher wishes to combine different data sources without knowing the exact link and merging rule. This paper considers different cartographic interpolation methods for interpolating attributes from German employment agency districts to German counties and vice versa. In particular, we apply dasymetric mapping as an alternative to simple area weighting both of which are based on estimated intersection areas. We also present conditions under which the choice of interpolation method does not matter and confirm the theoretical results with a simulation study. Our application to German administrative data suggests robustness of estimation results of interpolated attributes with respect to the choice of interpolation method. We provide weighting matrices for regional data sources of the two largest German data producers.

Keywords: area interpolation, dasymetric mapping, German administrative data
JEL: C49, C89, R10

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1.1 Introduction

With growing interest in research on the effects of recent German labour market reforms, researchers from both economics and social sciences alike have been increasingly concerned with combining information from different administrative data sources. In particular, researchers intend to combine information collected by the Federal Statistical Bureau (Statistisches Bundesamt) which is coded at the level of German counties with data from the Federal Employment Agency (Bundesagentur für Arbeit) which is reported for employment agency districts. Yet, the two sets of regions are geographically incompatible, i.e. one set of regions does not in general respect the boundaries of the other set and the two sets are not nested hierarchically. Hence, what is needed is an appropriate interpolation technique in order to transfer attributes from employment agency districts to counties and vice versa. The purpose of this paper is to provide a solution to this areal interpolation problem that may facilitate research based on data from both data sources in Germany.

In the geostatistics literature, interpolation techniques are always necessary if the spatial support of an attribute, i.e. its association with space, needs to be transformed to another spatial support. Depending on the type of spatial support involved, areas or points, the solution to this change of support problem (COSP) involves different spatial interpolation techniques.[2] If changing the support involves attributes that have been aggregated to a particular set of regions, interpolation methods have to solve the modifiable areal unit problem (MAUP). The MAUP arises because there is some arbitrariness in delineating regions and inference based on aggregated data depends on the level of aggregation and the grouping (zoning) of attributes (Openshaw and Taylor, 1981). Arbitrary zoning systems are often used by agencies for collecting and reporting socioeconomic data. Examples of such arbitrary zoning systems include German counties and employment agency districts in the case of the Federal Statistical Bureau and the Federal Employment Agency. Since the disaggregated data underlying the aggregated data are not available, a solution to changing the support from one of these zoning systems to the other (area-area COSP) necessitates a solution to the MAUP. In the literature, there are two distinct approaches to solving this problem: surface-oriented methods and cartographic methods.

[2]See Arbia (1989) and Gotway and Young (2002) for a description of the COSP and a review of different interpolation techniques.

Surface-oriented methods aim at a spatial smoothing of the aggregated data in order to re-construct the underlying disaggregated distribution of the attribute. For this purpose, Tobler (1979) proposes the so called smooth pycnophylactic interpolation. This method minimises curvature on the surface under the constraint that data from a source region can only be allocated to an intersecting target region ("pycnophylactic criterion"). Alternatively, inverse distance weighting uses a distance-decay weighting function for known values at centroids in order to interpolate attributes to unobserved positions (Bracken and Martin, 1989; Bracken, 1993, 1994). Kriging may also form part of the solution to the COSP by predicting attributes at unobserved positions based on an empirical semivariogram that gives the spatial autocorrelation between observed values (Cressie, 1993; Gotway and Young, 2002). All of these methods estimate a continuous surface representation from attributes that are known only for some source region and thus allow for calculating target area attributes by integrating this smoothed estimate over the target area. Since the resulting target values are based upon the underlying disaggregated distribution of the attribute, this approach solves the MAUP of the area-area COSP. Smoothing methods do, however, critically rely on knowing the central location of the source area and an adequate spreading function.

Cartographic approaches to the area-area COSP exploit information on the overlap of source and target areas to arrive at target area estimates. In the case of simple area weighting, attributes for target areas are estimated based on the intersection of source and target areas by assuming a uniform distribution of the attribute in the source zone (Goodchild and Lam, 1980). Simple area weighting thus solves the MAUP by reconstructing the underlying disaggregated distribution of the interpolated attribute based on a homogeneity assumption. Since this assumption is not always plausible, intelligent interpolation methods such as dasymetric mapping and statistical regression modelling have been proposed. Contrary to simple area weighting, these interpolation methods relax the homogeneity assumption by using auxiliary information on the source, target or some control zones, but still assume an even distribution within some subzones.

Dasymetric mapping or dasymetric weighting uses auxiliary information such as satellite images of populated and unpopulated areas to refine density estimates within the regions before allocating attributes to the target regions (Fisher and Langford, 1995). As an extension to this binary approach, it is also possible to distinguish more than two types of land use.

In this case, dasymetric mapping is only straightforward if the densities of different land use classes are known or somehow pre-defined (Eicher and Brewer, 2001). Alternatively, statistical regression modelling has been proposed to derive population density estimates for sub-regions by regressing the population of the source region on the different areas of land use (Flowerdew and Green, 1989; Langford et al., 1991; Yuan et al., 1997). Instead of using auxiliary information on the source or target region, Goodchild et al. (1993) develop a more general approach by using an external set of control zones for which uniform densities can be assumed. In a first step, control zone densities are estimated in a way similar to the procedure described by Flowerdew and Green (1989). In a second step, the estimated control zone densities are used to estimate target zone densities.

Approaches based on regression techniques all have to deal with a number of estimation issues such as the required non-negativity of estimated densities and the need to meet the pycnophylactic criterion. Moreover, count data such as population ("spatially extensive data") and proportional data such as average income or unemployment rates ("spatially intensive data") have to be treated differently (Goodchild and Lam, 1980). For spatially extensive data, Poisson regression has been proposed (Flowerdew and Green, 1989) as an alternative to constrained OLS regression (Judge and Yancey, 1986). In particular, Flowerdew and Green (1989) suggest an iterative Poisson regression using an EM algorithm to derive target area estimates. While this approach was first developed for spatially extensive data and binary auxiliary information only, extensions to continuous auxiliary data and spatially intensive data followed (Flowerdew and Green, 1992). Recently, Bayesian hierarchical models have been used to model Poisson responses with covariates that are spatially misaligned and thus unknown. Unlike the earlier approaches, the Bayesian approach allows for full inference of the distributions of estimated target zone attributes (Mugglin and Carlin, 1998; Mugglin et al., 2000; Best et al., 2000).

To sum up, ever more sophisticated methods have been applied to deal with the areal interpolation problem and to reduce the error involved in any interpolation exercise. Several authors have addressed the reliability of different methods and typically conclude that simple area weighting performs poorly compared with more sophisticated methods such as dasymetric mapping using regression frameworks (Goodchild et al., 1993; Fisher and Langford, 1995). Moreover, Fisher and Langford (1995) find some evidence that dasymetric approaches

may even yield better results than more elaborate methods based on statistical regression modelling. One reason for this latter finding may be that regression models fit global estimates that do not take into account the spatial heterogeneity in the link between the auxiliary information such as land-use type and the interpolated attribute. Thus, localised dasymetric mapping may perform better than more sophisticated methods and are typically easier to implement.

The purpose of this paper is to provide a solution to the areal interpolation problem between German counties and German employment agency districts. As a contribution to the areal interpolation literature, this paper considers an extended variant of dasymetric mapping that uses information on a local control variable that is available for both source and target region and that does not necessitate the use of regression techniques in order to arrive at refined estimates for target areas. We compare this dasymetric approach to simple area weighting as proposed by Goodchild and Lam (1980). Surprisingly, target area attributes for both interpolation methods are remarkably similar when interpolating data from German employment agency districts to German counties. We therefore introduce the concepts of local homogeneity and local similarity to explain this finding. Due to a high degree of local homogeneity and similarity, the choice of interpolation method does not have much influence on interpolated attributes for German counties. Thus, from a practitioner's point of view, even simple naive weighting seems a feasible solution in this specific case. A sensitivity analysis of the use of interpolated attributes as covariates in an economic analysis confirms that estimation results are not strongly affected by the choice of interpolation method.

The paper is structured as follows. Section 1.2 presents the framework for areal interpolation and suggests several interpolation methods. This section also derives conditions under which the choice of interpolation method does not have much influence on interpolated attributes. Using a Monte Carlo simulation, we demonstrate the effect of such conditions on the proposed methods of interpolation. Section 1.3 contains the application to German counties and employment agency districts including a sensitivity analysis. Section 1.4 summarises the main findings.

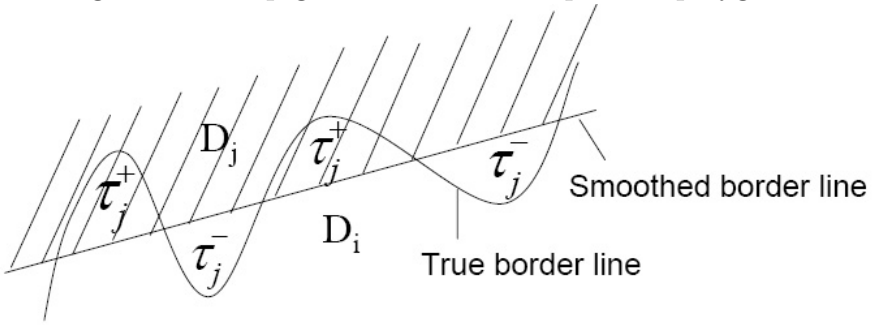
1.2 A framework for area interpolation

This section considers several cartographic areal interpolation methods for transferring attributes from a source to a target region and presents a simulation study in order to investigate the performance of the suggested methods in real world situations.

1.2.1 Polygon overlay and spurious polygons

Suppose we have two maps and each map contains a different disjoint regional classification of the same country. Denote $\{D_j\}_{j=1,\ldots,n}$ and $\{R_i\}_{i=1,\ldots,m}$ as two sequences of disjoint regions which are the elements of the two regional classifications. Let us denote μ as a measure of land area with the usual properties[3]. In an application the areas of intersection between D_j and R_i ($\mu(D_j \cap R_i)$) are typically unknown and have to be estimated by intersecting the two maps with a GIS procedure called vector polygon overlay based on the GIS software package ArcView. Some of the resulting intersection polygons are, however, likely to be spurious. Such spurious polygons result from polygon overlay if boundaries between the two regional layers are coincident in reality, but diverge due to digitising errors or differences in cartographic generalisation.[4] In this case - see Figure 1.1 - τ_j^+ is an area which is misleadingly assigned to region D_j. τ_i^- is a part of D_j which is misleadingly assigned to another region D_i. For simplicity, we assume absence of spurious polygons on the border of R_i.

Figure 1.1: Map generalisation and spurious polygons



In order to avoid such spurious polygons in the final map after polygon overlay, such slivers can, for example, be removed by either adding these slivers to the adjacent areas on

[3]See for example definition 4.1. in Elstrodt (1999).

[4]See Veregin (1989) and Chrisman (1989) for a discussion of error sources in map overlay. Recently, there has been some work on introducing general frameworks for error analysis in measurement-based GIS. See Goodchild (2004) for a motivation and Leung, Ma and Goodchild (2004) for a theoretical model.

a random basis or by connecting the end points of these slivers and dissolving the spurious polygon at both sides (Burrough, 1996). For the purpose of areal interpolation, spurious polygons have typically been neglected by erasing any entries in the intersection matrix below an arbitrary threshold. Due to the arbitrariness of this approach, we decided to keep spurious polygons since these spurious entries should leave the resulting interpolation unaffected as long as the induced measurement error is random and thus balances out. This should be the case if the error in digitised lines does not result from a misregistration that creates a uniform shift in the location of every point in the map, but instead is rather due to the uncertainty involved in hand digitising or differences in the degree of generalisation. Moreover, in most applications, the partitioning of the regions into sub-regions as a result of the intersection between D_i's and R_j's does not systematically depend on the topology of the border lines. In the real world this is because border lines between sub-regions are typically established on the basis of administrative considerations. For this reason, digitising errors, which are more severe when digitising complex curves than straight lines, should not introduce a systematic error component to the areal interpolation. For the exposition of the following framework of areal interpolation, we therefore assume that the induced measurement error is random and balances out, i.e. $\mu(\tau_j^- \cap R_i) \approx \mu(\tau_j^+ \cap R_i)$ for all i, j. This implies $\hat{\mu}(D_i \cap R_j) \approx \mu(D_i \cap R_j)$ with $\hat{\mu}(D_i \cap R_j)$ as an estimate of the true intersection $\mu(D_i \cap R_j)$. This also requires that there is only a minor misregistration error.

1.2.2 Cartographic interpolation methods

Without loss of generality, let us now consider different interpolation methods for transferring attributes from D_j, the source region, to R_i, the target region.[5] For this purpose, cartographic interpolation methods require weighting matrices that may be constructed based on the available estimates of areas $\hat{\mu}(D_j)$, $\hat{\mu}(R_i)$ and $\hat{\mu}(R_i \cap D_j)$. Note that there are two different kinds of attributes which have to be treated differently, i.e. for which different weighting matrices need to be used: frequencies (F), i.e. spatially extensive data, such as the number of job vacancies, participants in certain employment policies etc. and proportions (P), i.e. spatially intensive data, such as an unemployment rate.[6]

[5]Note that the opposite case is not considered but that our framework directly carries over.

[6]Goodchild and Lam (1980) introduced the terms spatially extensive data for frequencies and spatially intensive data for proportions in the context of areal interpolation.

Let us denote $f_{i,j}$ and $p_{i,j}$ as weights with the usual properties: $f_{i,j}$ and $p_{i,j} \geq 0$, $\sum_i f_{i,j} = 1$ and $\sum_j p_{i,j} = 1$ for all i, j. The general rule for interpolating data from D_j to R_i is

$$F_{R_i} = \sum_j F_{D_j} f_{i,j} \quad \text{for } i = 1, \ldots, n$$

where $f_{i,j}$ is an appropriate weight for frequency F_{D_j}, $j = 1, \ldots, m$ and

$$P_{R_i} = \sum_j P_{D_j} p_{i,j} \quad \text{for } i = 1, \ldots, n$$

where $p_{i,j}$ is an appropriate weight for proportion P_{D_j}, $j = 1, \ldots, m$. Depending on the underlying assumptions, there are several ways in which the weights $f_{i,j}$ and $p_{i,j}$ can be constructed. Apart from simple area weights, we focus here on two alternative approaches: naive binary weights[7] and some special form of dasymetric mapping (or dasymetric weighting) that refines the simple area weights by using an auxiliary variable such as the population density that is known for both source and target regions.

Naive binary weights

When interpolating proportions we allocate a weight of one to region D_j that shares the largest common area with R_i among all other intersecting regions. Thus,

$$p_{i,j} = \begin{cases} 1 & \text{if } \mu(R_i \cap D_j) = \sup_{D_l} \mu(R_i \cap D_l) \\ 0 & \text{otherwise} \end{cases}$$

for all i, j. These binary weights may be considered as a rule of thumb and can be obtained by simple visual inspection. They are a crude approximation and only yield reliable results if the attribute of the largest intersection area is representative for the entire target area. We include this naive binary weighting despite the much more sophisticated methods available because this rule of thumb is still being used by practitioners who are not familiar with the areal interpolation literature. Therefore, it is worthwhile to compare these weights to more sophisticated methods for our application to German counties and employment agency districts. Note that these binary weights cannot be used for the interpolation of frequencies because they typically violate the pycnophylactic constraint.

[7] These weights are also considered by Goodchild and Lam (1980), see their equation (13).

Naive weights can be estimated by replacing the true area sizes μ with their empirical counterparts $\hat{\mu}$, i.e.

$$\hat{p}_{i,j} = \begin{cases} 1 & \text{if } \hat{\mu}(R_i \cap D_j) = \sup_{D_l} \hat{\mu}(R_i \cap D_l) \\ 0 & \text{otherwise.} \end{cases} \tag{1.1}$$

for all i, j. This estimator yields the true $p_{i,j}$ if the ordering of the $\hat{\mu}(R_i \cap D_j)$ is same as for $\mu(R_i \cap D_j)$, i.e. differences in size of the intersection areas are greater than the measurement errors. If the maps are precise, this condition is likely to hold. Furthermore we assume for a given i that $\sup_{D_l} \hat{\mu}(R_i \cap D_l)$ is only attained for one D_l, i.e. for each R_i the intersection areas with D_l have different sizes if they are positive.

Simple area weighting

Following the proposed method by Goodchild and Lam (1980), simple area weighting assumes a homogeneous source zone density. Under this assumption, the simple area weights can be written as

$$f_{i,j} = \frac{\mu(R_i \cap D_j)}{\mu(D_j)} \quad \text{for all } i, j$$

in the case of frequencies and as

$$p_{i,j} = \frac{\mu(R_i \cap D_j)}{\mu(R_i)} \quad \text{for all } i, j$$

in the case of proportions using information on the intersection and area size of R_i and D_j only. Simple area weights can be estimated by replacing the true area sizes with their empirical counterparts:

$$\hat{f}_{i,j} = \frac{\hat{\mu}(R_i \cap D_j)}{\sum_i \hat{\mu}(R_i \cap D_j)}, \tag{1.2}$$

for all i, j and for $\hat{p}_{i,j}$ analogously. If all $\hat{\mu}(R_i \cap D_j)$ are almost their theoretical counterparts, it will likely give a precise estimate of the true value. In an application, however, $\hat{f}_{i,j}$ may be affected by the random measurement error of the map intersection.

Dasymetric mapping with local control variable

As a special form of dasymetric mapping, simple area weights can be refined by using a region-specific attribute that is known for source or target regions. Denote this control attribute as S_{R_i} and S_{D_j}. Under the assumption that the spatial distribution of this known

control attribute is highly positively correlated to the spatial distribution of the attribute to be interpolated to the target areas, one can use this information to re-estimate attribute densities of the intersection areas between source and target area. Thus, instead of assuming homogeneous source zone densities, we only assume a homogeneous density for intersection areas. In an application, one may use any known region-specific information that is highly positively correlated to the attributes to be interpolated. When using population, for example, S_{R_i} should be spatially intensive, i.e. $S_{R_i} = pop(R_i)/\mu(R_i)$ where $pop(R_i)$ is the number of individuals in $\mu(R_i)$. For frequencies we suggest

$$f_{i,j} = \frac{\mu(R_i \cap D_j)S_{R_i}}{\sum_i \mu(R_i \cap D_j)S_{R_i}} \quad \text{for all } i,j$$

with an appropriately defined S_{R_i}. For the merger of proportions we suggest

$$p_{i,j} = \frac{\mu(R_i \cap D_j)S_{D_j}}{\sum_j \mu(R_i \cap D_j)S_{D_j}} \quad \text{for all } i,j$$

with an appropriately defined S_{D_j}. These weights include the special case in which the region-specific variable does not contain any information, i.e. $S_{R_i} = S_R$ or $S_{D_j} = S_D$ for all i, j. In this case, the information is uniformly distributed across area space[8] and the weights simplify to the simple area weights by Goodchild and Lam (1980). Assuming S_{R_i} and S_{D_j} to be known numbers, dasymetric weights can be estimated similar to simple area weights by replacing the true area sizes with their empirical counterparts and, again, if spurious polygons are of minor importance, the estimates will be precise.

1.2.3 Misspecification of area interpolation

Area interpolation based on the proposed weighting schemes may not only be affected by the random measurement error induced by the polygon overlay. The construction of weights, i.e. interpolation method itself, may be misspecified if underlying assumptions do not hold. In particular, naive weights are misspecified unless there is a homogeneity of the attribute to be interpolated across all source zones that intersect with the target region. Somewhat less restrictive, simple area weighting assumes a uniform density distribution within the source region while dasymetric weights assume a uniform density distribution within the intersection areas. Clearly, none of the proposed interpolation methods needs to be appropriate if there

[8]For a given region i this requirement could be relaxed since it is only necessary that S_{R_i} does not vary in the neighbourhood of i.

is further local heterogeneity within the source or intersection regions. Among the proposed interpolation methods, dasymetric weighting should yield the least misspecified interpolation results. The question thus arises under which conditions misspecifications implied by naive binary weighting and simple area weighting result in large differences between the estimated frequencies F_{R_i} and proportions P_{R_i} across interpolation methods and under which conditions all methods yield very similar results. For this purpose, we introduce the concept of local homogeneity and global heterogeneity with respect to information S.

Definition 1: Local homogeneity with respect to information contained in S_i induces that $S_i \approx S_j$ for all i and all j in the direct neighbourhood of i.

Definition 2: Global $c-$heterogeneity corresponds to

$$sup_i \; inf_j \; |S_i - S_j| \leq c$$

for all regions i and all regions j in the direct neighbourhood of i and any $c \geq 0$.

It is then evident that a small c implies local homogeneity for all regions i. Having this in mind, it is easy to show that local homogeneity implies that simple area weighting and dasymetric weighting using the auxiliary information S yield very similar results.

Definition 3: Similarity of the regional entities R_i and D_j is defined by

$$sup_{R_i} |\mu(R_i) - sup_{D_j} \mu(R_i \cap D_j)| < \epsilon$$

for all i, j and any $\epsilon > 0$.

Similarity of the regional entities suggests that weights are similar across all weighting schemes. Clearly, if for all intersections i, j there is one large intersection that almost completely covers the reference region, differences between the interpolation methods tend to be small. In practice, a combination of local homogeneity and similarity of the two regional entities may yield very similar results for all interpolation methods and should also mitigate any remaining misspecification of the dasymetric approach.

1.2.4 Monte Carlo Evidence

It is interesting to examine the performance of naive and simple area weights as compared with dasymetric weights in the case of varying c-heterogeneity of the auxiliary information S and in case of varying similarity of regional entities.[9] For this reason, we perform a series of simulations for the prediction of proportion P_R. In order to make the simulation results comparable with our application in the following section, we use the same regional classification here for R and D. The number of sets R_i and D_j and the set of intersections is therefore identical to the empirical framework.[10] The remaining simulation framework is chosen as follows:

- We consider two different scenarios for the similarity of regional entities:
  - *a)* maximum dissimilarity of regional entities conditional on the set of intersections. This implies equal intersection areas for a given R_i, i.e. $\mu(R_i \cap D_j) = \mu(R_i \cap D_l)$ for all l s.t. $\mu(R_i \cap D_l) > 0$, while we randomly add a very small number to one of the intersection areas such that one intersection area is largest.
  - *b)* the intersection areas of our empirical framework. The two regional entities are similar and in many cases (80%) one single intersection area covers more than 95% of the target area.
- $P_D \sim U[5, 25]$ is an independently drawn random variable from a uniform distribution, i.e. no autocorrelation in P_{D_j}.
- S_D is drawn according to three different designs of spatial autocorrelation:
  - *i)* $S_D = 1$, i.e. no variation in the region-specific information.
  - *ii)* S_D is drawn element by element from $U[5, 25]$. If there is already a S_D assigned to the direct neighbourhood of S_{D_i} we compute $S_{D_i} = 0.2\epsilon_{D_i} + \bar{S}_{D_i}$, where $\epsilon_D \sim U[-10, 10]$ and $\bar{S}_{D_i}$ is the average over all neighbouring and already assigned S_{D_i}. This simulation design induces a weak positive spatial autocorrelation.[11]

[9] See also Fisher and Langford (1995) for an extensive Monte Carlo study for the comparison of different weighting schemes using data derived from a part of Leicestershire, England.

[10] In fact, we restrict the simulations to the case of western Germany. There are 327 target regions R_i and 635 source regions D_i. For further details see the application in the subsequent section.

[11] A Moran's I statistic confirms significant clustering of similar values of the region-specific information S_{D_i}.

– *iii*) $S_D \sim U[5, 25]$, random variation in the region-specific information.

Simulation designs *i-iii* allow for evaluating the relevance of the information S_D in an application. While we also vary the degree of similarity of regional entities, we keep the distribution of P_D constant over all designs. Since P_D is drawn independently, we consider here a framework without spatial autocorrelation in the source variable. In the case of spatial autocorrelation in the regional data that needs to be interpolated (P_D) and a very high degree of similarity of the regional entities, all three interpolation methods produce similar results.[12] Simulation results are presented in Table 1.1 where we relate the interpolated target variable, denoted by $\hat{P}_R$, using naive or simple area weights to the benchmark interpolation of P_R using dasymetric weights. We assume that the latter is also the true value of the target variable. Any biases and higher moments of the distribution are therefore due to the misspecification of the weighting schemes compared with dasymetric weighting.

As expected, Table 1.1 confirms that naive binary weighting performs worst in our simulation framework. It has the greatest bias and the largest standard deviation in all cases. This suggests that similarity of regional entities alone does not suffice to make naive weighting a comparably reliable alternative. We also observe that ignoring region-specific control variables in the case of simple area weighting biases results and the variance increases slightly (see ii) and iii)), especially if regional entities are dissimilar (see a)). Moreover, the misspecification is more severe in the case of a random variation in S (iii) than in the case of spatial autocorrelation (ii), i.e. less c-heterogeneity. In the case of similar regional entities and spatial autocorrelation of S (see b)ii)), the bias almost disappears. As expected, it is less important to include the region-specific information in this case. In the case of random variation in S, however, a higher degree of similarity between regional entities does not appear to have much of an influence (see b)iii)).

We conclude that without any precise information on the spatial distribution of the data and the degree of similarity of the regional entities, there is no way to tell how strongly research results are affected by the choice of interpolation method. In empirical applications, a sensitivity analysis may be useful to investigate the robustness of research results based on different interpolation approaches. Simulations in Fisher and Langford (1995) suggest that a dasymetric method gives better estimates than simple area weighting. This is in line with

[12]These cases are not presented but results are available on request.

Table 1.1: Monte Carlo evidence for the distribution of $(\hat{P}_R - P_R)/P_R$

	a) maximum dissimilarity			*b) case of Germany*		
	Mean	Sd	MSE	Mean	Sd	MSE[†]
Simulation i - uniform distribution of S						
Naive weights	0.0558	0.3917	0.1565	−0.0149	0.2287	0.0525
Area weights[‡]	0	0	0	0	0	0
Simulation ii - positive autocorrelation of S						
Naive weights	0.0619	0.3974	0.1618	−0.0167	0.2270	0.0518
Area weights[‡]	0.0047	0.0328	0.0011	−0.0005	0.0370	0.0014
Simulation iii - random distribution of S						
Naive weights	0.0597	0.4018	0.1650	−0.0103	0.2414	0.0584
Area weights[‡]	0.0022	0.0550	0.0030	0.0045	0.0589	0.0035

[†]Mean squared error

[‡]Area weights refer to simple area weighting with $S_{D_i} = 1$.

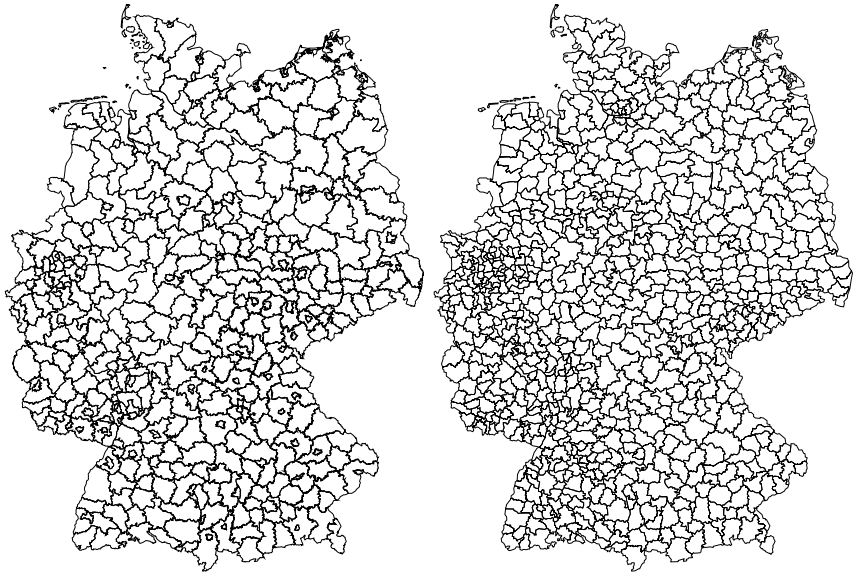
our simulation results. Due to a high degree of similarity, however, we expect differences between area weighting and dasymetric mapping in our application to Germany to be rather small.

1.3 Empirical application

The purpose of the empirical application is to identify an appropriate cartographic interpolation method in order to transfer attributes from German employment agency districts to German counties. As has been discussed in the introduction, these administrative agencies report data for different zoning systems which are spatially incompatible. In particular, both agencies provide important data for researchers in labour economics, other fields of economics and social sciences alike. Typically, German micro data include the county but not the employment agency district location of the individual or the household observation. Since current research on German labour market reforms often necessitates combining both data sources, solving this areal interpolation problem is thus of some importance and urgency.

Figure 1.2 shows a map of German counties (Kreise) and a map of Federal Employment Agency districts (Arbeitsamtsdienststellen). Think of the German counties as the R_i target regions with $i = 1, \ldots, 440$ disjoint entities. The Federal Employment Agency districts correspond to the D_j source regions with $j = 1, \ldots, 840$. We estimate county areas R_i and district areas D_j using the software package ArcView.

Figure 1.2: The German counties (left) and the German Federal Employment Agency districts (right)

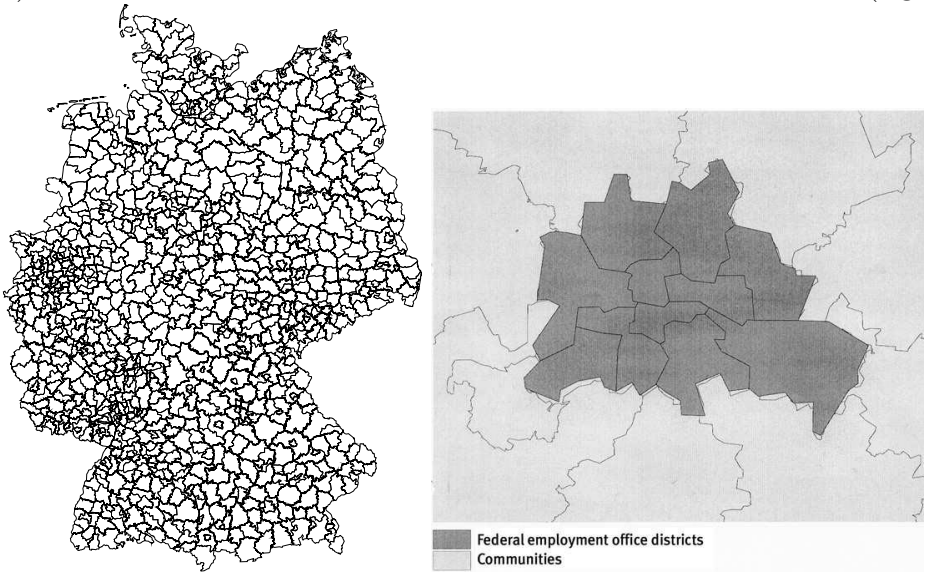


Since intersection areas that form the basis of any cartographic interpolation method are not readily available for German counties and German employment agency districts, we estimate county areas R_i, district areas D_j and their intersections $\hat{\mu}(R_i \cap D_j)$ using the GIS procedure of polygon overlay provided in the software package ArcView. Since the map of employment agency districts comes with a stronger generalisation than the map of German counties, intersecting both maps by polygon overlay results in spurious polygons, i.e. nonzero entries in the weighting matrix that are spurious due to digitising errors and the degree of generalisation.[13] Figure 1.3 to the left shows the resulting map from intersecting counties and districts. This intersection results in more than $3,600$ subregions, some of which are certainly spurious due to the measurement errors involved in intersecting the two maps. Spurious polygons can be seen at the border line of the Berlin area (see Figure

[13]In our particular case, map D was not available electronically. We therefore scanned the map in a raster data format. The raster data was then converted to vector data by digitising. Thus, in addition to smoothing errors due to cartographic generalisation, digitising errors may be another source of measurement error and spurious polygons.

1.3 to the right). In addition to a random measurement error due to digitising errors or cartographic generalisation, this figure suggests that there might also be a minor non-random misregistration error. Since the spurious entries in the intersection matrix seem to be trivial, however, even the slight misregistration error should not strongly affect the interpolation results.[14]

Figure 1.3: The intersection of German counties and German Federal Employment Agency districts (left) and stochastic measurement error at the Berlin border lines (right)



Based on these area estimates, we compute the three interpolation methods proposed in the previous section. More specifically, we perform a sensitivity analysis in order to test the robustness of estimation results with regard to the choice of interpolation method. This sensitivity analysis is of practical importance for researchers in Germany who work with micro data coded at the county level, but want to include regional socioeconomic data released by the Federal Employment Agency in their analysis.

We conduct a sensitivity analysis of the effect of certain regional labour market characteristics on the job-finding hazard of unemployed individuals in western Germany (excluding the Berlin area) between 1981 and 1997. We run the same estimation three times, each time including an area covariate that has been interpolated by one of the three proposed

[14]For German counties, we compared area estimates $\hat{\mu}(R_i)$ with their exact area size $\mu(R_i)$ which is officially released by the Federal Statistical Bureau (Statistische Ämter des Bundes und der Länder, 1999). On average, the measurement error induced by generalisation and digitising error is less than 0.1% of the true area size. This may be considered some evidence that the measurement error is negligible, even though for the intersections the measurement error should be larger.

methods, and discuss the results in light of the above theoretical considerations. The micro data set used for the analysis is the IAB[15] employment subsample 1981-1997 - regional file (IABS-REG) which is described in detail in Bender et al. (2000). The data set contains daily register data of about 500,000 individuals in western Germany with information on their employment spells as well as on spells during which they received unemployment compensation transfers from the Federal Employment Agency. The data set is a representative sample of employment that is subject to social security contributions and excludes, for example, civil servants and self-employed individuals. All individual information is coded at the level of the so called microcensus regions. These regional sub-divisions lump together up to four counties. There are 270 microcensus regions in western Germany. Based on this data set, we want to test the effect of two regional labour market indicators, namely the unemployment rate (P_{D_j}) and the ratio of unemployed individuals to vacancies in the region (F_{D_j}[16]) on the job-finding hazard of unemployed individuals. Both indicators are proxies for labour market tightness and may be expected to have a significant negative effect on the job-finding hazard of unemployed individuals in western Germany. More importantly, since these regional indicators are reported for employment agency regions only, they need to be interpolated to microcensus regions. Employment agency regions lump together three to four employment agency districts. Therefore, we can use the estimated intersections of employment agency districts and counties for an interpolation between the 270 microcensus regions and the 141 employment agency regions by aggregating the intersections to the level of microcensus and employment agency regions. Intersecting these two regional entities yields a total of 1,149 sub-regions.

Based on these intersections, we interpolate the two regional attributes, the unemployment rate and the ratio of unemployed individuals to vacancies, for the three proposed interpolation methods. As discussed in section 1.2, there are two possible reasons why estimated weights might not differ substantially between the alternative weighting schemes. First of all, there may be a high degree of local homogeneity in the region-specific information that is used for the dasymetric weighting approach. Here, we use regional labour force

[15]Institut für Arbeitsmarkt- und Berufsforschung in Nuremberg

[16]Instead of interpolating the ratio, we interpolate the number of unemployed individuals and vacancies as frequency data before calculating the ratio based on these interpolated attributes. As an advantage, this also allows for using the naive weight because the ratio of the interpolated frequencies equals the interpolation of the ratio and thus does not violate the pycnophylactic property.

densities as the region-specific information S because the distribution of the labour force should be highly positively correlated to other labour-market related attributes. Using a Moran's I statistic[17], we find evidence in favour of positive spatial autocorrelation, i.e. areas with high (low) labour force densities tend to be close to other regions with high (low) densities. Apparently, there is a high degree of local homogeneity or a low level of c-heterogeneity in the underlying region-specific control variable S. As a consequence, differences between area and dasymetric weighting should be rather small. Note also that the intersected spatial frameworks do show a high degree of similarity. In fact, almost 80% of all microcensus regions have an areal overlap of more than 95% with only one employment agency region (see also Figure 1.2). Thus, in most cases microcensus regions only have minor intersections with additional employment agency regions.

Due to a high degree of similarity between the two spatial frameworks and a high degree of local homogeneity, differences between simple area weighting and dasymetric weighting may expected to be rather negligible. Indeed, we find that the resulting weights do not differ substantially on average. In fact, with an average value that differs only in the 10th decimal place, dasymetric weights show an extremely similar distribution to simple area weights that assume a uniform distribution of the region-specific information. Standard deviations, percentiles as well as minima and maxima are also quite similar. For some subregions for which there is a low degree of local homogeneity within the neighbouring area, however, weights differ substantially across interpolation method. Table 1.2 looks at an extreme example - the Bremen metropolitan area - to demonstrate this point. Bremen is a large city in the north of Germany with about 500,000 residents and a relatively high labour force density compared to the surrounding rural areas (Diepholz, Wesermarsch, Osterholz, Rotenburg, Verden). Thus, while around 31 % of the area of the Bremen employment agency region intersects with the microcensus region of the same name, taking account of the fact that most of the labour force of the employment agency region works in this intersecting area results in a weight of almost 84 %.

We conclude that, on average, dasymetric and simple area weights do not differ sub-

[17] We calculate Moran's I using different weights for the spatially lagged vector based on the grid position. Using a weight of one for regions within a 0.4 degree radius of the grid location of the county, we get a test statistic of 0.21 ($z = 5.3$). Using a 0.8 degree radius the test statistic falls to 0.16 ($z = 8.9$) but again is highly significant. 0.1 degree correspond to 11.1 km along the longitude and between 6.5 to 7.5 km along the latitude.

Table 1.2: Weights $\hat{f}_{i,j}$ for the Bremen employment agency region for three interpolation methods

Employment agency region	Microcensus region	Area	Dasymetric	Naive
Bremen	Bremen	0.313	0.838	1
Bremen	Diepholz	0.002	0.000	0
Bremen	Wesermarsch	0.014	0.002	0
Bremen	Osterholz	0.652	0.157	1
Bremen	Rotenburg	0.016	0.002	0
Bremen	Verden	0.003	0.001	0

stantially because of a high degree of similarity between the two spatial frameworks and a high degree of local homogeneity. However, the choice of interpolation method may have an important influence for some selective regions with a high degree of heterogeneity in the auxiliary information within the local neighbourhood. We therefore look at two different samples for the sensitivity analysis, a full and a selective sample. The full sample includes all 255,100 unemployment spells[18] generated by 126,189 individuals and beginning between 1981 and 1997 in any microcensus region in western Germany.[19] The selective sample includes only unemployment spells from those microcensus regions whose estimated weighting schemes differed substantially.[20] Given the above results, we expect the analysis based on the full sample to be less sensitive with respect to the chosen interpolation method than the heterogeneous subsample. Even for the selective sample, however, estimation results may be quite robust across interpolation methods if the regional data to be converted, F_{D_j} and P_{D_J},

[18]Periods of registered unemployment cannot be identified easily given the data structure of the IAB employment subsample because there are gaps in an individual employment record whenever the individual neither works in a socially insured employment nor receives unemployment compensation from the Federal Employment Agency. As a consequence, labour market states such as self-employment, being out of labour force or being unemployment without being a benefit recipient are unobserved and indistinguishable (see Fitzenberger and Wilke, 2004). We therefore use a proxy for registered unemployment that is closely related to the receipt of unemployment compensation. Interruptions of these transfer payments may not exceed four weeks (six weeks in he case of a suspension period). Moreover, transfer receipt has to start within 10 weeks after the end of employment. If there is a gap of more than 12 weeks after the receipt of transfers, the unemployment spell is treated as censored to ensure that unobserved labour market states are not considered as an unemployment period.

[19]The sample has been restricted to individuals aged 18-52 at the beginning of the unemployment spell.

[20]A microcensus region belongs to the selective sample if the absolute deviation between the simple and the dasymetric weights is above the 99th or below the 1st percentile for either spatially intensive or spatially extensive data.

does not vary significantly between adjacent and nearby regions. In this case, even naive weights may produce reliable interpolation results. Indeed, a Moran's I statistic for both regional indicators finds significant spatial clustering of similar values.[21] As a consequence, even for a selective sample of regions for which weighting schemes differ significantly, interpolated unemployment rates (P_{R_i}) and unemployment-vacancy ratios (F_{R_i}) might be quite similar across interpolation methods.

Table 1.3: Summary statistics of interpolated unemployment rate $\hat{P}_D$ and u/v ratio $\hat{F}_D$[a] for the full and the selective sample by interpolation method

	Unemployment rate			*Unemployment-vacancy ratio*		
Method	Obs.	Mean	Std. dev.	Obs.	Mean	Std. dev.
Full Sample						
Naive weights	270	7.873	2.864	270	8.544	5.484
Area weights	270	7.864	2.833	270	8.371	5.121
Dasymetric weights	270	7.890	2.848	270	8.372	5.109
Selective Sample						
Naive weights	14	9.227	3.809	14	11.549	8.178
Area weights	14	9.226	3.646	14	9.833	6.788
Dasymetric weights	14	9.245	3.679	14	10.064	6.625

[a] The unemployment-vacancy ratio is treated as a spatially extensive attribute by interpolating the number of unemployed and the number of vacancies separately before calculating the unemployment-vacancy ratio.

Indeed, summary statistics of $\hat{P}_{R_i}$ and $\hat{F}_{R_i}$ at the level of microcensus regions in Table 1.3 confirm that differences between interpolation methods are levelled out. Even for the selective sample of 14 microcensus regions for which weights differed the most (see footnote 20), there is not much variation across the interpolated attributes. There is some more variation in the selective sample for the unemployment-vacancy ratio than for the unemployment rate, a result that is probably due to the fact that the auxiliary information used for the dasymetric weight of spatially extensive data, S_{R_i}, refers to employment agency regions. Since there

[21] See footnote 11 for details on the test statistic. Using a weight of one for regions within a 0.4 degree radius of the grid location of the county, we get a test statistic of 0.85 ($z = 16.3$). Using a 0.8 degree radius the test statistic is 0.72 ($z = 31.4$) which again is highly significant.

are fewer employment agency regions than microcensus regions, the corresponding dasymetric weighting should yield a coarser refinement of source zone densities than in the case of interpolating spatially intensive data. Still, summary statistics in Table 1.3 suggest that the choice of interpolation method even for the selective sample should not strongly affect regression results for these interpolated attributes.

For the sensitivity analysis, we estimate a proportional hazard model where the baseline hazard includes common fixed effects for individuals in the same labour market region.[22] This may be estimated using Cox's partial likelihood estimator (Cox, 1972). Including location-specific fixed effects in this estimator removes a potential bias of individual and labour market related variables that may result from omitting important regional labour market characteristics (Kalbfleisch and Prentice, 1980; Ridder and Tunali, 1999). In addition to the location-specific fixed effects we also take account of the fact that some individuals have repeated unemployment spells. Thus, we use the modified sandwich variance estimator to correct for dependence at the level of the individual (Lin and Wei, 1989).

Table 1.4 summarises estimation results for the unemployment rate and the unemployment-vacancy ratio for the full sample and the three interpolation methods. We control for education, sex, age, marital status, occupational status, economic sector, a set of year dummies as well as some indicators of prior employment history including total previous unemployment duration, tenure in the previous job and an indicator variable of whether there has ever been a recall from the previous employer. Summary statistics and estimation results using the full and the selective sample can be found in Appendix A[23].

As expected from the discussion above, the effect of the unemployment rate and the unemployment-vacancy ratio on the job finding hazard is extremely robust across the different interpolation methods for the full and the selective sample. In our empirical application, the effects of interpolated attributes on the estimated hazard ratios do not differ up to the 4th decimal place for the full and up to the 3rd decimal place for the selective sample. This even holds for naive binary weighting.

[22] We use labour market regions instead of microcensus regions because labour market regions are likely to be the relevant regional context in which individuals mainly seek employment. There are a total of 180 West-German labour market regions.

[23] Since estimation results across the various specifications are very similar, Appendix A only includes detailed results for the Cox model using the unemployment rate as a region-specific covariate. Moreover the estimation results only show the case of merging the unemployment rate based on a uniform distribution of the region-specific information.

Table 1.4: Cox PH model estimates for regional indicators by interpolation method and sample, IABS-REG, 1981-1997

	Full Sample		Selective Sample	
Merging Scheme	**Haz. Rat.**	**Std. Err.**	**Haz. Rat.**	**Std. Err.**
Unemployment-vacancy ratio				
Naive weights	0.989**	0.000	0.987**	0.001
Area weights	0.989**	0.000	0.986**	0.001
Dasymetric weights	0.989**	0.000	0.985**	0.001
Unemployment rate				
Naive weights	0.967**	0.001	0.976**	0.005
Area weights	0.967**	0.001	0.971**	0.005
Dasymetric weights	0.966**	0.001	0.973**	0.005

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

We conclude that in the specific case of interpolating data between German employment agency regions and microcensus regions the choice of interpolation method does not substantially affect our estimation results. It even seems safe to take the simplest approach available to the researcher: an interpolation based on simple binary weights. Due to a high degree of local homogeneity in S, a high degree of similarity of the regional entities and a strong positive spatial autocorrelation of the data to be interpolated, this is likely to be a result that is unique to this particular application. Hence, researchers applying the above approach to a different set of regional entities should be aware that these factors have an important effect on the robustness of their results. Also, they should check the degree of spatial autocorrelation of the spatially misaligned data. If there is spatial clustering of dissimilar values, interpolation is likely to be much more sensitive to the choice of interpolation method than in our particular application. Therefore, researchers are advised to examine the conditions of local homogeneity, similarity of regional entities and positive or negative spatial autocorrelation in detail before choosing an interpolation method. If there is evidence that even the dasymetric weighting approach may be seriously misspecified and no positive spatial autocorrelation of the attributes to be interpolated mitigates this misspecification, other more sophisticated methods might be necessary to arrive at satisfactory results.

1.4 Conclusion

This paper presents several cartographic methods for interpolating spatially misaligned data from German employment agency districts to German counties. We compare interpolation results from naive binary weighting, simple area weighting and a more sophisticated dasymetric weighting approach that makes use of auxiliary regional information. In particular, we use a control variable that is available for both spatial frameworks and that is highly positively correlated to the spatially misaligned data in order to refine the source zone density estimates. Dasymetric weighting thus necessitates less strong assumption concerning the spatial distribution of the the spatially misaligned data than simple area weighting to solve the modifiable areal unit problem (MAUP).

Furthermore, we identify conditions under which all interpolation methods including naive binary weighting yield comparable and reliable results. Under a high degree of local homogeneity in the auxiliary information used for the dasymetric weighting approach and under a high degree of similarity between the two regional classifications, the choice of interpolation method does not matter. We confirm these theoretical results with a simulation study. As a sensitivity analysis for the area interpolation between employment agency regions and microcensus regions, we compare the effects of interpolated attributes on the job-finding hazard of unemployed individuals using all three interpolation methods. Our application suggests robustness of estimation results with respect to the choice of interpolation method. In addition to a high degree of local homogeneity in the auxiliary information and a high degree of similarity between microcensus and employment agency regions, local homogeneity in the attribute to be interpolated further mitigates any differences between the three methods. We conclude that in our particular application even the naive binary weighting yields reliable results. Moreover, even in the case that the dasymetric approach does not solve the MAUP, i.e. the underlying assumptions do not hold, the remaining misspecification should be of minor importance due to the spatial autocorrelation of the attributes that need to be interpolated. Our interpolation techniques thus allow for combining data from the two largest German data producers, the Federal Employment Agency and the Federal Statistical Bureau and thus provides an important preparatory work for the empirical analyses in chapter (2) and (5). The estimated weighting schemes are freely accessible to the research community and can be downloaded from $ftp://ftp.zew.de/pub/zew-docs/div/arntz-wilke-weights.xls$.

Appendix

A - Summary statistics for the full and the selective sample of unemployment spells, IABS-REG, 1981-1997

	Full Sample		Selective Sample	
	Mean	**SE**	**Mean**	**SE**
Unemployment duration (in days)	293.24	443.86	285.55	407.46
Female	0.41	0.49	0.44	0.50
Married	0.46	0.50	0.44	0.50
Married female	0.21	0.41	0.21	0.41
Age < 21	0.08	0.28	0.07	0.26
Age 21-25	0.23	0.42	0.21	0.41
Age 31-35	0.14	0.35	0.14	0.35
Age 36-40	0.11	0.31	0.13	0.32
Age 41-45	0.10	0.30	0.11	0.31
Age 46-49	0.07	0.26	0.08	0.27
Age 50-53	0.08	0.27	0.08	0.27
Low education	0.38	0.49	0.36	0.48
Higher education	0.04	0.20	0.05	0.23
Low education x female	0.16	0.37	0.17	0.37
Higher education x female	0.02	0.13	0.02	0.15
Apprenticeship	0.07	0.25	0.06	0.25
Low skilled worker	0.34	0.48	0.32	0.47
White collar worker	0.25	0.43	0.30	0.46
Part-time work	0.08	0.27	0.09	0.28
Agriculture	0.03	0.17	0.02	0.13
Inv. goods industry	0.20	0.40	0.17	0.38
Cons. goods industry	0.12	0.32	0.08	0.28
Construction	0.15	0.36	0.12	0.33
Services	0.31	0.46	0.38	0.49
Tenure in previous job (in months)	27.20	38.09	26.88	38.64
Previous recall	0.06	0.23	0.05	0.22
Total unemp. duration (in months)	8.43	15.03	8.29	14.70
1983-1987	0.32	0.47	0.32	0.47
1988-1991	0.19	0.39	0.19	0.39
1992-1997	0.34	0.47	0.34	0.47
Unemployment rate[a]	9.70	3.38	9.34	3.56
Number of spells	255,100		83,104	
Number of individuals	126,189		24,674	
% right-censored	28.4		29.7	

[a] Regional information has been merged using the uniform distribution of the region-specific information $S_{D_j} = 1$.

Cox PH model estimates (hazard ratios) using the full and the selective sample, IABS-REG, 1981-1997

	Full Sample		Selective Sample	
Variable	**HR**	**(SE)**	**HR**	**(SE)**
Female	1.112**	(0.011)	1.127**	(0.035)
Married	1.219**	(0.008)	1.227**	(0.031)
Married female	0.539**	(0.013)	0.583**	(0.022)
Age < 21	1.217**	(0.010)	1.281**	(0.045)
Age 21-25	1.103**	(0.008)	1.141**	(0.029)
Age 31-35	0.985†	(0.009)	1.001†	(0.029)
Age 36-40	1.000	(0.011)	1.002	(0.031)
Age 41-45	1.001	(0.011)	1.011	(0.035)
Age 46-49	0.968*	(0.013)	0.930*	(0.037)
Age 50-53	0.831**	(0.015)	0.823**	(0.037)
Low education	0.883**	(0.009)	0.847**	(0.023)
Higher education	0.792**	(0.020)	0.779**	(0.044)
Low education x female	0.968*	(0.013)	1.035*	(0.041)
Higher education x female	1.149**	(0.030)	1.120**	(0.089)
Apprenticeship	1.082**	(0.013)	1.136**	(0.045)
Low skilled worker	0.798**	(0.009)	0.845**	(0.023)
White collar worker	0.752**	(0.010)	0.805**	(0.023)
Parttime work	0.806**	(0.016)	0.829**	(0.035)
Agriculture	1.317**	(0.020)	1.333**	(0.092)
Inv. goods industry	0.927**	(0.010)	0.926**	(0.027)
Cons. goods industry	0.925**	(0.011)	1.029**	(0.036)
Construction	1.221**	(0.010)	1.347**	(0.041)
Services	0.984†	(0.009)	1.024†	(0.025)
Tenure in previous job	0.995**	(0.000)	0.994**	(0.000)
Previous recall	0.781**	(0.012)	0.756**	(0.031)
Total unemp. duration	0.995**	(0.000)	0.997**	(0.001)
1983-1987	1.245**	(0.008)	1.252**	(0.033)
1988-1991	1.332**	(0.009)	1.365**	(0.039)
1992-1997	1.085**	(0.009)	1.122**	(0.032)
Unemployment rate[a]	0.967**	(0.001)	0.971**	(0.005)
Log-likelihood	-1,217,399.365		-121,004	
$\chi^2_{(30)}$	18,724.774		1,961.91	

Significance levels : † : 10% * : 5% ** : 1%

[a] Using the merging scheme with $S_{D_j} = 1$.

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Chapter 2

The Geographical Mobility of Unemployed Workers. Evidence from Western Germany

Abstract

Using a competing risks framework of exiting unemployment to jobs in a local or a distant labour market area, this paper investigates whether unemployed individuals in western Germany choose search strategies that favour migration away from depressed regions. Empirical results are obtained from a stratified Cox proportional hazard model that allows for region-specific fixed effects. This approach reduces biases due to unobserved regional heterogeneity that affects both the migratory behaviour of jobseekers and the regional conditions of interest. Estimates based on this stratified approach indicate that at least skilled men are responsive to local labour market conditions and are more likely to leave regions with unfavourable re-employment opportunities. Without region-specific fixed effects, no such evidence can be found. The implied downward bias from omitting region-specific effects is consistent with the idea that pleasant living conditions compensate for higher regional unemployment.

Keywords: migration, unemployment duration, competing risks, unobserved regional heterogeneity
JEL: J61, J64, R23

2.1 Introduction

It has often been argued that interregional mobility plays a crucial role in equilibrating regional employment and wage disparities. The underlying notion is that unemployed workers leave depressed regions in order to find employment in regions that offer better employment prospects. For the US, Blanchard and Katz (1992) find this adjustment mechanism to be quite effective. In European countries including Germany, however, interregional labour mobility lags behind mobility levels in the US, Canada, Japan and Australia (Eichengreen 1991, Braunerhjelm et al., 2000). More importantly, even though unemployment and wage differences are important factors in determining migration in Europe, recent findings suggest that the elasticities of aggregate migration flows with respect to unemployment and wage differentials are lower than in the US (Puhani, 2001). Decressin and Fatàs (1995) examine regional adjustment dynamics in Europe and find that the adjustment process following an adverse regional employment shock takes much longer than in the US. A study by Möller (1995) also suggests that interregional mobility in Germany is an important means of equilibrating regional employment disparities in the long run, but that it takes 2-3 years before interregional mobility begins to respond to an adverse regional employment shock.

The effectiveness of migration as an equilibrating mechanism ultimately depends on the migratory decisions of individual people. In particular, given the high level of unemployment in Germany, the willingness and ability of unemployed workers to seek employment in more prosperous regions and to migrate away from depressed areas is of central concern if migration is to contribute to reducing regional disparities. Recent empirical evidence on internal migration in western Germany is rather mixed. Decressin (1994) looks at migration flows between states in western Germany and finds that these flows tend to go from high to low unemployment regions. In contrast, a recent study by Windzio (2004) on the determinants of individual mobility between the southern and northern states in western Germany even suggests that individuals in high unemployment regions have lower migration probabilities. These mixed results are not untypical when compared to international findings. While Herzog and Schlottmann (1984) for the US and Tervo (2000) for Finland find evidence that high regional unemployment encourages individuals to migrate out of the region, UK studies by Pissarides and Wadsworth (1989) and Hughes and McCormick (1994) suggest that high regional unemployment levels even discouraged mobility during the 1970s and 1980s.

One explanation for these mixed results lies in the duration dependence of the mobility decision. As suggested by Goss and Schoening (1984), not controlling for search duration may bias estimation results due to unobserved heterogeneity of migration probabilities across search time. Therefore, recent research explicitly models the migratory behaviour adopted by unemployed individuals within a hazard model specification of unemployment durations. By distinguishing between the competing risks of exiting unemployment to different regional labour markets, this approach provides information on the actual search strategy of unemployed workers. The few existing studies, however, show contradictory results. Using a Gompertz proportional hazard model with gamma distributed unobserved individual heterogeneity, the Finnish study by Kettunen (2002), for example, does not indicate any significant effect of local labour demand on the migration hazard, i.e. the hazard of finding employment via residential mobility. Based on a Cox proportional hazard analysis, the US study by Yankow (2002) finds that higher employment and wage levels significantly reduce the migration hazard, while the unemployment rate and regional employment growth do not exert any significant influence. One reason for this contradictory evidence might be that neither study takes account of unobserved regional heterogeneity. If the unobserved heterogeneity in the attractiveness of a region as a place for living, i.e. its level of regional amenities affects both the migratory behaviour of jobseekers as well as regional labour market conditions, omitting such factors may seriously bias estimation results. According to the compensating differentials theory, individuals in regions with unpleasant living conditions must be compensated in some way for the implied disutility - either by lower unemployment rates, lower rents or higher wage levels (Roback, 1982; Marston, 1985; Elhorst, 2003). Thus, unemployment and wage differentials may reflect individuals' underlying preferences for certain regions.[1] Several studies confirm these compensating differentials in the USA where workers are willing to accept lower wages in regions offering more attractive living conditions (Blomquist et al., 1988; Gyourko and Tracy, 1991; Greenwood et al., 1991). The evidence for compensation of unemployment differentials is less clear-cut. While some US studies do not indicate that local amenities contribute to unemployment differentials (Vedder and Gallaway, 1996; Partridge and Rickman, 1997), some recent European studies confirm the expected relationship

[1] At a given point in time, the cross-sectional variation of unemployment and wage levels reflect both equilibrium and disequilibrium forces. Amenity differentials are considered an equilibrium force that results in persistent regional wage and employment disparities.

(Aragon et al., 2003; Lopez-Bazo et al., 2005). If labour market conditions are related to the attractiveness of a region as a place for living and also affect the migratory behaviour of jobseekers, omitting such factors in a model of migratory behaviour may bias estimation results.

The purpose of this study, therefore, is to contribute to the literature by re-examining the migratory behaviour of unemployed job-seekers in the context of western Germany using a competing risks hazard model framework that takes account of unobserved heterogeneity at the regional level. In particular, I estimate a stratified partial likelihood estimator (Ridder and Tunali, 1999) that introduces region-specific fixed effects to the Cox partial likelihood estimator (Cox, 1972). This approach mitigates biases due to omitted regional characteristics that may be correlated with the variables of interest such as the level of regional amenities. The inclusion of region-specific fixed effects is possible because the analysis is based on the employment subsample 1981-1997 - regional file of the IAB (*Institut für Arbeitsmarkt- und Berufsforschung*). This register data set is well-suited to the proposed analysis because the sample contains a sufficient number of relatively rare interregional mobility events and thus facilitates the analysis of the migratory behaviour of unemployed individuals. More importantly, the data set allows mobility events to be observed over a long time period providing sufficient time variation to identify the model with region-specific fixed effects. The proposed analysis thus allows the responsiveness of search strategies to be re-examined using less restrictive assumptions than earlier studies by Kettunen (2002) and Yankow (2002). In particular, my aim is to investigate whether unemployed workers in western Germany adopt search strategies that favour migration away from regions with unfavourable re-employment opportunities. I also look at the responsiveness to local labour market conditions among different labour market segments such as low-skilled and skilled individuals and take account of some passive and active labour market programs that may affect individual mobility decisions.

The findings of this study indicate that controlling for unobserved heterogeneity at the regional level influences estimation results. While there is significant evidence that men experience higher migration hazards in regions with unfavourable job-finding conditions when using a stratified partial likelihood estimator, there is no such evidence when using an unstratified estimator that is subject to unobserved regional heterogeneity. The implied down-

ward bias from omitting region-specific effects is consistent with the idea that pleasant living conditions compensate for unfavourable job-finding conditions. Looking at several labour market segments reveals that low-skilled men are less responsive to local labour market conditions than their skilled counterparts. Moreover, unemployment benefits seem to reduce mobility among all individuals, while active labour market measures only exert a minor regional locking-in effect on women.

The outline of the paper is as follows. The next section introduces a model of job search across space. This job search model predicts that various region-specific factors, including job-finding conditions as well as regional amenities, will affect the job search strategy of unemployed jobseekers. The institutional background, data and methodological approach are discussed in section 2.3. Section 2.4 presents estimation results. Section 2.5 concludes.

2.2 A search model with search across space

The theoretical framework adapts the two-sector framework by Fallick (1992) to the case of job search across two regional labour markets by introducing mobility costs between regions. In this stationary framework, individuals search simultaneously[2] across a local labour market l and a distant labour market d. The framework can be described as follows:

- Jobseekers are risk-neutral and maximise the expected present value of job search V^u, discounted over an infinite horizon at rate r.

- Real wage offers from each labour market k (k=l,d) are drawn from known distributions $f_k(w)$ and arrive according to a Poisson process with job offer rate $\alpha_k\sigma(e_k)$. α_k captures the exogenous job offer conditions in the labour market and $\sigma(e_k)$ is an increasing and concave function of the endogenously determined search effort devoted to labour market k. This framework implies that the instantaneous probability of receiving more than one job offer at a time is zero.

- Searching the two labour markets comes at cost $c_k(e_k)$ which is an increasing and convex function of the search effort in market k. As an assumption, the returns to

[2]This is a generalisation of the systematic search literature that considers the job searcher to sequentially sample regions, firms or sectors according to the expected returns from searching on these sub-markets (see Salop, 1973; McCall and McCall, 1987).

searching labour market k exceed the search costs and thus there is always a positive amount of job search in each labour market.

- The jobseeker additively derives utility from income and regional amenities a_k. In the context of this paper, regional amenities apply to a rather broad concept of natural amenities (e.g. climate, landscape), consumer amenities (e.g. theatres, diversity of consumption possibilities) as well as housing prices and cost-of-living differentials that affect the welfare of an individual living in a particular region. Recent research has shown that regional amenities significantly affect migration decisions and should thus be included in a job search framework across space (see Hunt and Mueller (2004) and chapter (5) of this dissertation). The framework thus allows for amenity differentials across space, an extension of the search model that has also been discussed by Damm and Rosholm (2003).

- Accepting a job offer from the distant labour market entails flow-type and lump-sum type moving costs. Flow-type moving costs pc_d capture the permanent utility loss (e.g. psychological costs, loss of social capital) related to the residential move, while lump-sum moving costs mc_d cover the once-only costs of relocation. By assumption, moving costs always exceed utility differences of unemployed job search in the local and the distant labour market so that jobseekers never move to d due to amenity or cost-of-living differentials that may also affect the real value of transfer receipt. As a consequence, the framework only allows for contracted migration, a restriction that can be justified by the marginal relevance of speculative migration in European labour markets (Molho, 1986).

According to this framework, the unemployed jobseeker chooses the reservation wage w_k^r and the allocation of search effort across k that maximises the expected present value of continuing job search in the local area V_l^u:

$$\begin{aligned} rV_l^u &= b_l - c_l(e_l) - c_d(e_d) + a_l \\ &\quad + \alpha_l \sigma(e_l) \int_{w_l^r}^{w_{max}} (V_l^e(w) - V_l^u) dF_l(w) \\ &\quad + \alpha_d \sigma(e_d) \int_{w_d^r}^{w_{max}} (V_d^e(w) - V_l^u) dF_d(w) \end{aligned} \tag{2.1}$$

This flow value consists of the instantaneous utility of being unemployed locally (i.e. the real value of transfer payments b_l and local amenities a_l minus search costs), the expected surplus of a local job multiplied by the probability of receiving a job offer locally and the expected surplus of a distant job that involves interregional residential mobility multiplied by the probability of receiving a job offer in this market. Assuming, for simplicity's sake, that a worker keeps his job forever, the value of employment in the local market $V_l^e(w)$ may be written as

$$rV_l^e(w) \;=\; w + a_l \tag{2.2}$$

with r as the discount rate. The corresponding value of employment in region d is given as

$$rV_d^e(w) \;=\; w + a_d - pc_d - r\,mc_d. \tag{2.3}$$

with $r\,mc_d$ as the flow value of the lump-sum moving cost. The unemployed jobseeker accepts a job offer from region k if the value of being employed at the offered wage equals or exceeds the value of continuing job search in the local area. Noting that $r(V_d^e(w) - V_l^u) = w + a_d - pc_d - r\ mc_d - r\ V_l^u$, the reservation wage for the distant labour market that equates $V_d^e(w_l^r)$ and V_l^u is

$$w_d^r \;=\; r\ V_l^u - a_d + pc_d + r\ mc_d. \tag{2.4}$$

The corresponding reservation wage for the local region is

$$\begin{aligned} w_l^r \;&=\; r\ V_l^u - a_l \qquad (2.5)\\ &=\; b_l - c_l(e_l) - c_d(e_d)\\ &\quad + \frac{\alpha_l \sigma(e_l)}{r} \int_{w_l^r}^{w_{max}} (w - w_l^r) dF_l(w)\\ &\quad + \frac{\alpha_d \sigma(e_d)}{r} \int_{w_d^r}^{w_{max}} (w - w_d^r) dF_d(w) \end{aligned}$$

Note that the reservation wage in the distant region may be expressed in terms of the local reservation wage:

$$w_d^r \;=\; w_l^r + (a_l - a_d) + pc_d + r\ mc_d. \tag{2.6}$$

Thus, w_d^r diverges from w_l^r in order to compensate the job mover for the relocation costs and the amenity differential between both regions. It follows that individuals with high

permanent relocation costs due, for example, to strong family ties to the local region, are less likely to accept a job offer from a labour market that involves migration than others. The reservation wage for the distant region must also compensate for the flow value of the lump-sum moving cost. However, because an infinite time horizon is assumed for the jobseeker, this effect is relatively minor. In a search model with a limited time horizon, the effect of the relocation costs may instead be severe. Kettunen (1994) has shown that in the presence of moving costs, the reservation wage for the distant labour market increases dramatically as the end of the working life approaches.[3] As a consequence, the hazard of moving to the distant region may be close to zero for individuals who enter unemployment during their final years in the labour force. In the subsequent empirical analysis, such age effects thus need to be taken into account.

Comparative statics in Appendix A indicate that reservation wages for both local and distant jobs increase with improving job offer arrival rates and a shift of the wage offer distribution to the right anywhere in the economy. Both reservation wages also rise with transfer payments and decrease with higher search costs. By contrast, higher relocation costs and changes in amenities affect both reservation wages differently, as can easily be seen from equation (2.6). While higher relocation costs reduce the local reservation wage due to the reduced choosiness of the individual, the reservation wage for the distant labour market increases since higher costs have to be compensated by higher wages. The increase is less than unity, however, because the individual is less picky anywhere in the economy. The same qualitative results hold for both types of relocation cost. The magnitude of the effect is much smaller, however, in the case of the once-only costs mc_d. Finally, an increasing level of local amenities reduces the local reservation wage and increases the reservation wage for the distant region.

Besides determining the reservation wages for both markets, the job searcher endogenously allocates search effort across the two labour markets. Optimal search effort e_k^* is derived by differentiating V_l^u with respect to e_k

$$c_k'(e_k^*) \;=\; \alpha_k \sigma'(e_k^*) \int_{w_k^r}^{w_{max}} (w - w_k^r) dF_k(w) \quad \forall k = d, l \tag{2.7}$$

[3]The opposite is the case in a finite-horizon search model without any relocation costs (Gronau, 1971). The reservation wage decreases and the hazard of leaving unemployment increases because the value of continuing search decreases.

Thus, optimal search effort for any regional labour market equates the marginal benefit with the marginal costs of searching this market. Comparative statics in Appendix B indicate that deteriorating conditions in one region shift search effort towards the other region. In particular, a higher local job-offer arrival rate, an improvement in the local wage offer distribution and a higher level of local amenities raise the local search effort while reducing search effort in the distant region. More generous transfer payments reduce search efforts in all labour markets. By contrast, higher relocation costs unambiguously raises the local and reduces the distant search effort.

The probability that an individual i with characteristics x who is unemployed at the beginning of period t makes a transition to employment in k during this period is now given by the probability of being offered a job in k and the probability of accepting it:

$$h_k(t|x_i) \quad = \quad \alpha_k(e_k^*|x_i) * [1 - F_k(w_k^r(x_i))]$$

From the above framework, it follows that the local employment hazard $h_l(t|x_i)$ and the migration hazard $h_d(t|x_i)$ depend on conditions in all regions by affecting the search strategy of the jobseeker. While the effect of changing labour market conditions in the distant region on the migration hazard reflects both the direct effect that stems from changing exogenous conditions and the induced effects of a changing search strategy[4], changes in local labour market conditions should affect the migration hazard only via the induced changes of the search strategy. A two-region model thus allows for some inferences on the search strategy of unemployed jobseekers. In a one-region model such inference is not possible because the direct effect on the exit hazard and the induced indirect effect cannot be separated. For a detailed discussion see also Fallick (1993). The two-region framework thus allows for testing whether unemployed jobseekers are sensitive to labour market conditions and adjust their search strategies accordingly. This responsiveness to labour market conditions is highly desirable if labour mobility is to contribute to reducing regional disparities. Empirically testing whether local labour market conditions have the expected and desired effect on the migration hazard may be difficult though, if there are unobserved factors such as regional

[4]A higher job offer arrival rate α_d or a rightward shift of the wage offer distribution μ_d should attract search effort to the distant region and increase reservation wages in both markets. The positive effect of improved exogenous conditions and higher search efforts on the migration hazard thus contrast a negative reservation wage effect. The direction of the effect is thus theoretically ambiguous. See Mortensen (1986) for a detailed discussion. However, the stronger the shift of search effort to the distant region, the more likely should be a positive effect on the migration hazard.

amenities that may affect both an individual's search strategy and the observed interregional wage and employment differentials. Previous studies such as Kettunen (2002) and Yankow (2002) did not account for unobserved heterogeneity at the regional level. The following section therefore proposes an alternative approach that allows for region-specific fixed effects and that should thus be better suited to test the predictions of the above framework.

2.3 Data and Methodology

This section presents the data as well as the methodological approach and discusses the choice and definition of covariates used in the subsequent analysis.

The IABS-REG 1981-1997 The analysis is based on the IAB employment subsample 1981-1997 - regional file (IABS-REG) which is described in detail in Bender et al. (2000). The IABS-REG contains spell information on a 1 % sample of the population working in jobs that are subject to social insurance contributions. As a consequence, the sample does not represent self-employed individuals and life-time civil servants. For western Germany, the sample includes spell information on about 500,000 individuals for whom employment histories can be reconstructed on a daily basis including the microcensus region of the workplace. In addition, the data contains spell information on periods for which the individual received unemployment compensation from the Federal Employment Agency (*Bundesagentur für Arbeit*). During the period of analysis, unemployment compensation (UC) in Germany consisted of unemployment benefits UB (*Arbeitslosengeld*) and unemployment assistance UA (*Arbeitslosenhilfe*). The former is an insurance benefit with limited eligibility. Whether someone is entitled to receive UB and the maximum entitlement length depends on previous job tenure and age. By contrast, the less generous unemployment assistance is tax-funded and is paid indefinitely if entitlements to unemployment benefits have been exhausted. Since unemployment assistance is means-tested, however, it only applies to those who lack other financial resources such as spouse income. As a consequence, there is a gap in the IABS-REG record whenever an individual continues to be unemployed after the exhaustion of unemployment benefits and does not receive any unemployment assistance. Since such a gap in the IABS-REG record is indistinguishable from other unobserved labour market states such as being out of the labour force or self-employed, it is necessary to define unemployment spells

according to a suitable bound (Fitzenberger and Wilke, 2004). In the following analysis, unemployment spells are defined according to a bound introduced by Lee and Wilke (2005) and which is shown in Figure 2.1:

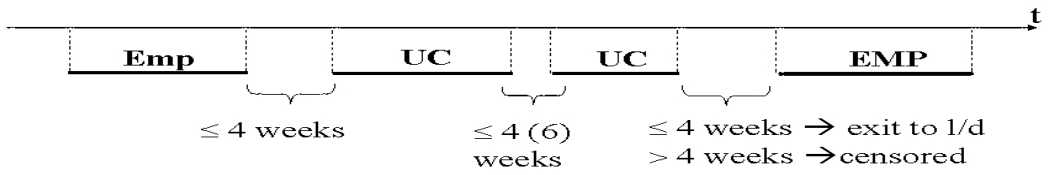

Figure 2.1: Bounding unemployment in the IABS-REG

This unemployment proxy[5] imposes the following restrictions:

- The receipt of transfer payments has to begin within four weeks after the end of employment (EMP). This restriction tends to exclude voluntary unemployed since voluntarily quitting a job results in a suspension period for the receipt of unemployment compensation of four to six weeks. Thus, individuals included in the analysis should mainly be displaced workers.[6]

- Gaps between periods of transfer receipt may not exceed four weeks. In the case of a suspension period due to rejecting an acceptable job offer, the gap must not be longer than six weeks. Otherwise, the unemployment spell is considered right-censored since individuals with long gaps are likely to have left unemployment for an unobserved labour market state.

- Unemployment spells are considered right-censored if the gap between the end of transfer receipt and the beginning of employment exceeds four weeks. This restriction tends to treat spells of long-term unemployed who are not entitled to receive unemployment assistance as censored, but at the same time censors spells of individuals who are no longer actively seeking employment. The resulting sample of unemployment spells therefore is not representative for unemployed people who are not eligible for

[5]For comparison, I also did the subsequent estimations for an alternative proxy of unemployment using all non-employment periods with an initial period of transfer receipt. Estimation results for the two definitions are very similar so that I show results only for this unemployment proxy.

[6]There is some imprecision in the exact length of the suspension period. Thus, there may still be some voluntary unemployed included in the analysis.

unemployment assistance because they have other sources of income, but should still represent the majority of unemployment spells in Germany.

Non-censored spells of unemployment exit to employment that is subject to social insurance contributions. Since participating in certain active labour market programs involves a socially-insured job (e.g. work creation schemes, long training measures), the exit to employment observed in the data may be either an exit to a regular or to such a subsidised job. The exit hazard therefore includes a program participation hazard which should be taken into account in the econometric analysis by controlling for the local availability of such programs.

Exits to employment may occur either locally or outside the local labour market. Since the IABS-REG includes information on the microcensus region of the workplace, interregional mobility can easily be identified.[7] I define interregional mobility as movements between non-adjacent labour market regions (*Arbeitsmarktregionen*). Labour market regions (LMRs) comprise typical daily commuting ranges such that for the majority of individuals the workplace is located within the LMR. Finding employment in a non-adjacent LMR, i.e. outside a 50 to 80 km radius, therefore usually necessitates residential mobility. In western Germany, there are 180 labour market regions (LMR) that lump together 326 counties.

I restrict the analysis to spells of unemployment starting between 1983 and 1995 in western Germany (excluding West Berlin). Excluding the latter two years ensures that for each spell of unemployment we observe at least two years before the end of the observation period. Earlier years are excluded due to missing information on major covariates. The sample includes individuals aged 26 to 60 years at the time of job loss. Applying the above unemployment definition, these restrictions yield a sample of 137,845 unemployment spells. 33.1% of these unemployment spells are right-censored, while 59.9% exit to a local job within the extended local labour market region and 7.0% exit to a job in a distant labour market region. Table 2.1 shows that exit types and the median unemployment duration strongly differ by sex and educational attainment. In particular, women remain unemployed longer than their male counterparts, a result that likely reflects the weaker labour force attachment among women. Another interesting finding from Table 2.1 is that individuals who lack any

[7]During the unemployment spell, the actual whereabouts of the individual are not known, however. This means that it is not possible to distinguish between contracted and speculative migration. Contracted migration as modelled by the theoretical framework is much more common than speculative migration, however (Molho, 1986).

vocational training have longer unemployment durations and are less mobile geographically. Given the high level of unemployment among low-skilled individuals, the responsiveness of this sub-group to local labour market conditions may be particularly important for the equilibrating role of migration. The following analysis thus runs separate estimations not only by sex, but also by skill-level.

Table 2.1: Unemployment duration and exit types by sub-samples, IABS-REG 1983-1995

	Men			Women		
	All	**LS**[a]	**S**[a]	**All**	**LS**[a]	**S**[a]
% exit to						
local	65.3	63.1	65.8	50.6	45.7	52.3
non-local	7.8	4.0	8.7	5.8	2.6	6.9
censored	27.0	32.9	25.6	43.7	51.7	40.8
Unemployment spells						
Median duration (days)	114	134	109	181	209	167
Number of spells	87,260	16,562	70,698	50,585	13,144	37,441

[a] LS - Low-skilled without vocational training; S - Skilled with vocational training or tertiary education

A stratified Cox proportional hazards model The econometric analysis focuses on two competing hazard rates, the hazard of finding a job within the extended LMR (h_l) and the hazard of finding a job in a distant LMR (h_d), i.e. the migration hazard, as a function of time spent in unemployment. Since the focus of the analysis is not on the shape of the hazard function, a competing risks form of the semi-parametric Cox proportional hazard model (Cox, 1972) is an appropriate choice for the proposed analysis. A clear advantage of the semi-parametric Cox estimator compared with parametric specifications is that the baseline hazard is specified fully flexible. This avoids any biases that result from misspecifying the shape of the baseline hazard in parametric specifications (Lancaster, 1990).

Assuming that the two competing risks are independent conditional on all covariates included in the model[8], the exit-specific hazard rate of the Cox proportional hazard model for individual i may be written as

$$h_k(t_i|x_i) \quad = \quad h_k(t_i)exp(x_i(t)\beta_k)$$

where t_i is the elapsed duration of unemployment for individual i, $h_k(t)$ is the exit-specific baseline hazard with $k = d, l$ and $x_i(t)$ is a vector of both time invariant and time-varying covariates. β_k is the vector of parameters of interest.

As a caveat of this approach, estimation results based on this specification may be biased due to unobserved individual and unobserved regional heterogeneity. To account for individual heterogeneity, I include a large number of individual covariates as described in detail in the next section. Moreover, in a model with a non-parametric and thus fully flexible baseline hazard, unobserved individual heterogeneity has been found to have little effect on the estimated coefficients (Meyer, 1990). In fact, when comparing the estimates to a parametric log-logistic accelerated failure time model with gamma distributed unobserved individual heterogeneity, no significant differences could be detected.[9] To account for unobserved regional heterogeneity, I modify the specification by estimating a fully flexible baseline hazard for each local labour market (LMR) j. This stratified Cox partial likelihood estimator (SPLE) removes any biases that result from unobserved, time-invariant characteristics of the local labour market region (LMR) such as unobserved regional amenities. A competing risks form of the SPLE may be written as:

$$h_{kj}(t_{ij}|x_{ij}, \nu_j) \quad = \quad h_{kj}(t_{ij}, \nu_j)exp(x_{ij}(t_{ij})\beta_k)$$

with t_{ij} as the duration of unemployment of individual i in LMR j. $h_{kj}(t_{ij}, \nu_j)$ is the baseline hazard in LMR j and is allowed to depend on an unobserved location-specific fixed effect ν_j.

[8]This is a critical assumption since estimation results will only be consistent estimates of the true parameters if this independence assumption holds. The latent durations for the two exits may, however, be dependent if there are unobserved factors that affect both exit hazards (Gordon, 2002). To account for relevant factors, the analysis includes a broad array of factors that affect general job-finding chances (see detailed discussion in the next section) as well as factors that affect migration probability (e.g. marital status, children). Unfortunately, there is no information on home ownership, an important aspect of the migration decision. Thus, I cannot rule out some stochastic dependence between both exit types, but stick to the approach to ease comparability with previous studies that also use independent competing risks without taking account of unobserved regional heterogeneity. Alternatively, generalised dependent risk models relax the assumption of independent competing risks (Gordon, 2002). Moreover, Honoré and Lleras-Muney (2006) describe methods to bound the effect of covariates in the case of dependent competing risks.

[9]Estimation results for the alternative parametric model are available from the author upon request.

This nuisance parameter along with the baseline hazard cancels out of the likelihood function. The possibility of removing stratum-specific fixed effects has already been discussed by Kalbfleisch and Prentice (1980) and Chamberlain (1985). Ridder and Tunali (1999) discuss the conditions under which such an approach is appropriate when using time-varying covariates. In particular, covariates have to be weakly exogenous, i.e. an explanatory variable x_t may not depend on observed exits from unemployment in the same labour market region in period $\tau \geq t$. Due to the discrete nature of the data, this exogeneity condition may be problematic for some contemporaneous regional indicators. I therefore use lagged variables for those regional indicators for which such an simultaneity issue is likely to arise (see Appendix C). Moreover, the inference is based on robust standard errors that take into account the clustering of individuals within labour market regions (see Lin and Wei, 1989). Otherwise, standard errors of covariates at the regional level may be biased downward (Moulton, 1990).

Individual-level covariates Following the theoretical framework, individual covariates should include factors that determine an individual's relocation costs. The empirical literature on mobility decisions typically finds individuals with local family ties, older, less-educated individuals and homeowners to have higher moving cost. The subsequent analysis thus includes marital status, children[10], educational attainment and age as important covariates. Controlling for age effects is particularly important given the zero migration hazard that is likely to occur when approaching the end of the working life (Kettunen, 1994). Unfortunately, the IABS-REG does not include information on housing tenure. On the other hand, the data structure of the IABS-REG allows the employment history of the unemployed jobseeker to be constructed which should capture some heterogeneity across individuals in terms of their job-finding chances and mobility costs. The previous job status and the previous sector of activity, for example, may capture different job-finding chances. Previous unemployment experiences are likely to reduce general re-employment chances due to a depreciation of human capital and possible stigma effects. Moreover, financial resources to bear moving costs may have been depleted during a long unemployment period, thus reducing the migration hazard. Lower migration hazards may also occur in the case of a long job tenure and thus residential immobility. Similarly, having been recalled from the previous

[10]The information on whether someone has children is likely to contain some measurement error as employers do not always report the current household context of their employees reliably.

employer may increase a person's local attachment due to waiting for another future recall. The analysis also includes previous wage income to capture heterogeneity in an individual's productivity level that is not reflected by their formal education. Since the literature on the returns to migration typically finds a positive selection of migrants with regard to ability, an individual with higher previous wage income for a given educational attainment is likely to have higher returns to migration and should thus have higher migration propensities (Greenwood, 1997).

Another set of indicators included in the analysis concerns the unemployment compensation system. As discussed previously, individuals may either receive unemployment benefits (UB) or the less generous but indefinitely paid unemployment assistance (UA). Whether someone is entitled to receive UB and the maximum entitlement length depends on previous job tenure and age. During the 1980s, maximum entitlements became more generous, especially for older unemployed (see Appendix E). Moreover, the income replacement rate, i.e. the share of the previous wage income that is substituted for by unemployment compensation, has increased successively from 60% to 68% for unemployment benefits and from 50% to 58% for unemployment assistance. According to the theoretical framework, more generous transfer payments - either in terms of a longer entitlement period for unemployment benefits or higher replacement rates - should reduce search intensities and reservation wages everywhere. Consistent with these notions, Hunt (1995) finds evidence that the increasing generosity of unemployment compensation in Germany throughout the 1980s increased the duration of unemployment. Moreover, more generous unemployment benefits may also be a means of postponing migration decisions which would reduce the migration hazard (Hassler et al., 2005). Including the income replacement rate and the maximum entitlement period[11] is thus meant to control for these institutional changes throughout the study period.

Regional-level covariates According to the theoretical framework, jobseekers are expected to shift search effort towards other regions if local labour market conditions are bad relative to conditions in other regions. Since the observation period encompasses almost 15 years, however, the same relative labour demand conditions, for example, may occur for

[11]The maximum entitlement period is computed from the IAB-REG information on previous job tenure within a 7-year period prior to the benefit claim and age. This estimate is likely to be an upper bound of the true entitlement length because previous unemployment periods may partially have exhausted these entitlements.

different levels of aggregate labour demand. Deteriorating overall employment prospects, however, should affect both exit hazards similarly and leave the allocation of search effort between the local and the distant labour market unaffected. In order to separate the effect of aggregate conditions and relative conditions between local and distant regions, I therefore choose the following type of specification for capturing the effect of labour demand conditions X_{LM}:

$$\beta_k X_{LM} \quad = \quad \beta_{k1} U_l/V_l + \beta_{k2} \frac{U_l/V_l}{U_d/V_d} \tag{2.8}$$

with $k = d, l$ and U_k/V_k as the unemployment-vacancy ratio (**uv-ratio**) in region k, and $\frac{U_l/V_l}{U_d/V_d}$ as the **relative uv-ratio**.[12] Due to this specification, β_{k1} reflects the effect of deteriorating overall labour demand conditions because relative labour demand conditions are held constant, i.e. $\frac{\partial h_k}{\partial U_l/V_l}|_{\frac{U_l/V_l}{U_d/V_d}}$. Moreover, changing relative labour demand conditions while holding the local labour demand conditions constant, i.e. $\frac{\partial h_k}{\frac{U_l/V_l}{U_d/V_d}}|_{U_l/V_l}$, reflect changing conditions in the distant labour market. Following the discussion on the possible inference with respect to the search strategy in a two-region model in section 2.2, β_{d2} thus captures both the direct effect of changing labour market conditions and the indirect effect of a changing search strategy. By contrast, β_{l2} allows for some inference on the search strategy as changing conditions in distant regions should affect h_l mainly[13] via a response of the search strategy. In particular, the theoretical framework suggests that improving labour demand conditions in the distant region (while keeping local conditions constant) raise reservation wages in all labour markets and result in a shift of search effort from the local to the distant region. A deteriorating relative uv-ratio should thus result in a decreasing local exit hazard. In addition, we should also observe an increasing migration hazard if the positive direct effect of improved job offer conditions plus the positive effect from a shift of search effort to the distant region outweigh the negative reservation wage effect.

In addition to local job-finding conditions, the theoretical framework suggests that income prospects should be another major concern for those searching for a job. Similar to equation (2.8), the analysis includes the local **wage level** and the **relative wage level**, i.e.

[12] Unfortunately, neither the uv-ratio nor the unemployment rate are available at a high level of regional disaggregation for different skill groups. Thus, only an average regional indicator for all sub-segments of the labour force can be used.

[13] As an underlying assumption, conditions in both regions need to be independent. Otherwise changes in the labour market conditions of one region may partially reflect changing conditions in the other region.

the expected local wage level for individual i relative to the expected aggregate wage for this individual.[14] in order to capture the effects of both relative wage conditions and different wage levels on the transitions to local and non-local employment. Unfortunately, regional cost-of living indices are not available for Germany so that wages have been deflated by the aggregate German Price Index (1985 = 100) which is released by the Federal Statistical Bureau. Since interregional wage disparities should be related to unobserved interregional disparities in cost-of-living and amenities, the included wage measures are likely to be biased. Similarly, estimates for relative labour demand conditions may be biased if amenities compensate for unfavourable job-finding conditions as discussed in the introduction. Although both rent and amenity conditions may have changed during the period under study, including fixed effects should capture average differences between the regions and should thus mitigate such biases. In addition, changes in the **population density** may to some extent reflect changing amenity and rent conditions as an increasing population density may come with additional consumer amenities as well as with disamenities that result from a lack of housing space and higher housing prices.

Furthermore, I use several indicators to further characterise local labour market conditions which may have an impact on transitions to either local or non-local jobs. For these local labour market conditions, significant impacts on the migration hazard indicate a changing search strategy in response to these local conditions, while effects on the local employment hazard should always be a mixture of such induced effects and the direct effect of changing local conditions. One important covariate included in the following analysis is the local **accommodation ratio**, i.e. the ratio of individuals in work creation schemes to individuals who are either unemployed or participating in such programs.[15] Not including such a measure could bias the estimated effect of local labour demand conditions because

[14]Wages have been estimated separately for each region and for the economy as a whole based on the IAB-R01, a data set that is closely related to the IABS-REG (see Appendix C). Estimates from Mincer-type wage regressions have been used to predict wage levels for each individual in the sample in his home region and in the economy as a whole.

[15]In Germany, active labour market programs (ALMP) mainly consist of either public employment schemes or training measures with an emphasis on the latter. In 1997, almost 270,000 people entered training programs, while around 75,000 people entered work creation schemes in western Germany (Caliendo et al., 2003). Unfortunately, a time series encompassing the years between 1983 and 1995 is only available for work creation schemes. The effect of training measures cannot therefore be tested. However, since regions with extensive work creation also tend to offer extensive training measures, the accommodation ratio may also proxy for the general accommodation of ALMP.

such programs typically try to cushion adverse labour market conditions and exits to regular employment and to subsidised employment in the context of such labour market programs are indistinguishable in the IABS-REG. Moreover, an extensive local supply of work creation schemes may reduce interregional mobility.[16] Such regional locking-in effects may occur if joining such programs substitutes for regular non-local employment. In this case, the extensive availability of such programs may reduce search efforts in the regular labour market, especially in distant regions. The attractiveness of local versus distant job search may also depend on whether job-finding conditions tend to improve or worsen. For this reason, **local employment growth** in individual *i's* skill group is used to capture such effects. Job-finding conditions may also be affected by employment concentrations in particular sectors. Sectors may differ, for example, by their cyclical sensitivity which may affect local job-finding chances. In order to capture such effects, I include **employment shares by sector**, distinguishing between agriculture, the consumption goods industry, the investment goods industry, construction, services and retail. In order to control for interregional differences in the educational composition of the workforce, I include a measure of the **employment share of individuals with a tertiary education**. The average quality of the workforce may affect local job-finding prospects for jobseekers of different educational background depending on the production technology and the degree of competition between skill groups. Conditions that may affect local job-finding chances may also be captured by the **share of male unemployed**. A high share of male unemployed typically prevails in regions with declining male-dominated industries and may thus reflect an additional congestion effect on the labour market for men that is not captured by the uv-ratio. Finally, I control for the size of the labour market by including the **employment level**. If there were increasing returns to job matching in large markets, the local employment hazard should positively depend on the labour market size (Petrongolo, 2001).

Finally, annual fixed effects are included to measure aggregate cyclical effects and capture possible changes that are due to the re-unification of West and East Germany in 1990. A detailed description of the data, its sources, definitions and summary statistics are shown in Appendix C. All regional indicators are coded at the level of labour market regions which

[16] Fredriksson and Johansson (2003) and Lindgren and Westerlund (2003) present evidence that mobility is lower among Swedish participants of ALMP. Moreover, Westerlund (1997, 1998) also finds evidence that an increasing number of participants in ALMP reduces flows of out migration in Sweden.

are likely to be the most relevant for the job search behaviour of unemployed jobseekers. Moreover, in order to ensure comparability across estimates, all regional covariates have been standardised at the level of labour market regions, i.e. an increase of a covariate by one refers to an increase by one standard deviation compared with the average labour market region in the period 1983-1997.

Marginal effects on interregional mobility As discussed previously, the estimated coefficients are of direct interest when it comes to testing the predictions of the theoretical model that job searchers adjust their spatial search strategies in response to changing labour market conditions. These coefficients do not say anything, however, about the qualitative or quantitative effect of the covariates on the probability of leaving unemployment via migration. This is because, in an independent competing risk model, the likelihood of exiting via a specific type of exit depends on covariate estimates for all exit-specific risks (Thomas, 1996). In particular, the probability of an unemployed person with characteristics x leaving unemployment for a job in a distant labour market, i.e. the cumulative migration probability is given by

$$\Pi_d(t|x) = \int_0^t h_d(t|x)S(t|x)dt$$

with $h_d(t|x)$ as the migration hazard and $S(t|x)$ as the overall survival function. Thus, this probability is also a function of the covariate parameter for the local employment hazard so that the qualitative effect on migration probability may even be positive if the estimated effect of covariate x_i is negative for both h_d and h_l. In addition to the estimated coefficients, the following analysis thus also looks at the marginal effect of a covariate on $\Pi_d(x)$:

$$\kappa_k(t|x) = \frac{\partial \Pi_k(t|x)}{\partial x_i}$$

I simulate these marginal effects by calculating the difference between the probability of exiting via migration within two years $\Pi_k(730|\bar{x})$ for a reference worker with average individual and regional characteristics and the respective probability that is due to a marginal change of a covariate[17]. Due to the stratification technique, I obtain separate simulated marginal

[17]For all continuous variables, I simulate the marginal effect of increasing x_i by a standard deviation. For all dummy variables the effect refers to the difference between zero and one. Thus, note that the marginal effect for dummy variables does not exactly refer to the base probability $\Pi_k(730|\bar{x})$ that is included in the subsequent result tables.

effects for each local labour market region. I thus calculate the average marginal effect across all strata $\bar{\kappa}_k$ by averaging across all j labour market specific marginal effects κ_{k_j}. Based on 500 repetitions, I determine the 2.5% and the 97.5% percentiles of the conditional marginal effects bootstrap distribution.

2.4 Estimation Results

The first subsection discusses estimation results for men and women to detect major result patterns and to compare the results across specifications with and without unobserved regional heterogeneity. Next, we look at the heterogeneity of these effects across skill groups. Given the high level of unemployment among low-skilled individuals in Germany, the responsiveness of this labour market segment to labour market conditions may be particularly important for the equilibrating role of migration.

2.4.1 Mobility effects when accounting for unobserved regional heterogeneity

Tables 2.2 and 2.3 show estimation results for the local employment and migration hazard for males and females, respectively. Successively expanding the set of regional covariates did not significantly affect the findings so that results are shown only for the full set of covariates.[18] Each table contains estimated hazard ratios for both the unstratified and the stratified partial likelihood estimator. According to the clustering test statistic proposed by Ridder and Tunali (1999), the inclusion of labour-market specific strata is highly significant in all specifications. Moreover, taking into account unobserved regional heterogeneity affects the estimation results for some regional covariates and suggests serious biases in the case of unobserved regional heterogeneity. Thus, marginal effects are only displayed for the stratified model. Before looking at the effects of regional covariates in more detail, however, some individual level covariates also deserve some discussion. In fact, individual characteristics dominate unemployment experiences of unemployed jobseekers while labour market characteristics often have a comparatively small effect only.

[18] Estimation results for an expanding set of covariates are available from the author upon request.

Individual level covariates The family background of the jobseeker has quite different effects on the unemployment experiences of men and women. Married women have a significantly lower employment hazard both locally and non-locally than their male counterparts and thus experience longer unemployment periods. Similarly, with an increase in the cumulative migration probability of 27.7%, having children strongly enhances geographic mobility among men, but has the opposite effect on women (−36.0%). These gender differences are likely to reflect the lower labour force attachment of women as well as child-rearing responsibilities. Moreover, the mobility-enhancing effect of having children on the mobility of men may to some extent reflect weekly long-distance commuting which due to data limitations cannot be distinguished from residential movements.

The employment hazard for individuals above the age of 50 strongly decreases. The probability of leaving unemployment via migration even decreases by around 70% for men and women. This is in line with much lower exit rates at the end of the working life in the presence of high mobility costs (see Kettunen, 1994). In Germany, such an effect may have been reinforced by the institutional framework. During the observation period, older unemployed often were entitled to a very long receipt of unemployment benefits. These transfers in addition to compensation payments by their previous employers enabled older unemployment to retire early without really seeking a new employment (see Lüdemann et al., 2004). The extremely low exit rates for individuals with UB entitlements of more than two years are in line with such disincentive effects, especially for geographic mobility. In the case of men, UB entitlements that exceed two years significantly reduce the migration probability by 29.3% compared with 5.4% for local exits. Lower exit rates for individuals with short or no entitlements to unemployment benefits may thus seem at odds with the theoretical expectations. However, this is likely to reflect unobserved negative selection of this group because short or zero entitlement periods only occur if someone has not been working continuously during the last years. Similar result patterns for the effect of the length of UB entitlements on transitions to employment with the same data set have been found by Biewen and Wilke (2005). Moreover, in line with the theoretical expectations, a higher income replacement rate decreases the employment hazard for both men and women.

As expected from the literature on migration, abler individuals as reflected in higher

formal education and higher pre-unemployment wages[19], have higher probabilities of leaving unemployment via migration. Compared with individuals who have vocational training, having only a high school degree, for example, decreases the likelihood of migration for men (women) by 41.3% (38.0%) while a tertiary education increases the likelihood of being mobile by 30.5% (13.4%). Concerning the effect of the pre-unemployment wages, Biewen and Wilke (2004) find shorter unemployment durations among previously well-earning unemployed. The estimation results for the competing risks model suggest that this effect is due to a higher migration hazard rather than due to a higher local employment hazard of individuals in a higher wage quintile. This is in line with the migration literature that often finds a positive selection of abler migrants since these individuals are more likely to benefit from migration in terms of wage gains (Greenwood, 1997). On the other hand, the percentage of homeowners should be higher among previously well-earning unemployed which would rather suggest lower migration rates. The mobility-enhancing effect of high pre-unemployment wages, however, clearly dominates since the cumulative migration probability even increases by 36.3% for men and 42.9% women, respectively.

Previous unemployment periods increase the local employment hazard and leave the migration hazard unaffected. This might reflect lower reservation wages of previously unemployed individuals, but a lack of willingness to invest in geographic mobility. By contrast, long previous unemployment periods significantly decrease both the local employment and the migration hazard and thus prolong current unemployment. Moreover, prolonging previous unemployment by three months has a stronger effect on the migration hazard and thus reduces the cumulative migration probability by 2.7% for men. This may reflect a combination of exhausting financial resources and the discouraging effect of a precarious employment history on investing in geographic mobility. Finally, as expected, having ever been recalled from the previous employer approximately halves the cumulative migration probability but increases the local employment hazard.

[19]Of course, simultaneously controlling for previous wage income captures some differences between skill levels. Still, I decided to also include the previous wage income because it reflects some of the heterogeneity in ability and productivity that is not captured by the formal education.

Table 2.2: Estimated hazard ratios and marginal effects for the local employment hazard (h_l) and the migration hazard (h_d) for the unstratified (UPLE) and the stratified partial likelihood estimator (SPLE), IABS-REG 1983-1995, Men

	UPLE		**SPLE**			
	$exp(\hat{\beta}_l)$	$exp(\hat{\beta}_d)$	$exp(\hat{\beta}_l)$	$exp(\hat{\beta}_d)$	$\hat{\kappa}_l{}^{\dagger}$ (%)	$\hat{\kappa}_d{}^{\dagger}$ (%)
$\Pi_k(730\|\bar{x})(\%)$					60.2	3.6
Individual characteristics						
Married	1.223**	1.236**	1.217**	1.219**	10.2*	11.3*
Children	1.118**	1.337**	1.113**	1.356**	4.3*	27.7*
Age 26-29	1.097**	1.092*	1.093**	1.095*	4.5*	5.1
Age 30-34	1.045**	1.066	1.045**	1.068	2.0*	4.5
Age 40-44	0.966†	0.916†	0.961†	0.925†	-1.8	-5.6
Age 45-49	0.930**	0.849**	0.920**	0.865**	-3.8*	-9.9*
Age 50-54	0.883**	0.685**	0.877**	0.705**	-5.7*	-24.7*
Age $>$ 55	0.412**	0.201**	0.408**	0.208**	-38.7*	-69.9*
High school degree	0.913**	0.550**	0.909**	0.553**	-2.9*	-41.3*
Higher education	0.748**	1.253**	0.755**	1.254**	-21.1*	30.5*
Unskilled blue-collar	0.880**	0.886**	0.880**	0.884**	-6.1*	-6.2*
White-collar	0.615**	1.256**	0.620**	1.243**	-27.0*	44.3*
1st & 2nd wage quintile	0.818**	0.840**	0.821**	0.839**	-8.9*	-7.7*
4th & 5th wage quintile	0.988	1.366**	0.989	1.376**	-2.2*	36.3*
Prev. job tenure	0.983**	0.965**	0.983**	0.965**	-0.5*	-2.5*
Previously unemployed	1.280**	0.959	1.264**	0.960	14.8*	-12.0*
Total prev. unemp. dur.	0.984**	0.964**	0.984**	0.965**	-0.7*	-2.7*
Recall from prev. employer	1.485**	0.544**	1.444**	0.551**	22.2*	-52.7*
Max. UB 0 months	0.544**	0.708**	0.555**	0.704**	-29.4*	-12.5*
Max. UB 1-12 months	0.752**	1.101*	0.745**	1.079†	-15.9*	20.5*
Max. UB 13-24 months	1.003	0.985	1.001	0.964	0.1	-3.6

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... table 2.2 continued

Max. UB > 24 months	0.869**	0.682**	0.874**	0.658**	-5.4*	-29.3*
Income replacement rate	0.972**	0.943**	0.974**	0.943**	0.3*	-3.0*
Fourth quarter	0.500**	0.747**	0.513**	0.745**	-31.7*	-3.5
Year and sector dummies‡	X	X	X	X		
Regional covariates						
UV-ratio	0.887**	0.878**	0.882**	0.890**	-6.1*	-5.8*
Relative uv-ratio	0.967**	1.028	0.967*	1.112**	-2.4*	12.2*
Wage level	0.949**	1.048†	0.949**	1.052†	-3.0*	7.4*
Relative wage level	0.983†	0.932**	0.990	0.926**	-0.2	-6.6*
Employment growth	1.033**	0.968	1.030**	0.969	1.8*	-4.2*
Accommodation ratio	1.062**	1.073**	1.048**	1.032	2.5*	1.1
Male unemployment	0.919**	0.806**	0.914**	0.784**	-4.1*	-18.2*
Share of HS emp.	0.965†	0.995	1.030	0.932	2.0	-7.7*
Employment level	1.009	1.024†	1.126	1.010	6.2	-4.3
Population density	0.980†	0.994	0.945	0.955	-3.2	-2.4
Agriculture	1.015	0.971	1.072*	1.151	3.1*	11.1*
Investment goods ind.	1.084*	0.984	1.100†	1.161	4.5	10.8
Consumption goods ind.	1.103**	0.981	0.959	0.861	-2.0	-12.2*
Construction	1.096**	1.079**	1.098**	1.118	4.6*	7.0
Services	1.099**	1.040	1.155**	1.088	7.2*	1.6
Number of spells	87,260	87,260	87,260	87,260		
Number of exits	56,952	6,782	56,952	6,782		
Log-likelihood	-592,263.7	-68,049.7	315,481.1	-36,176.2		
$\chi^2(df)$ clustering test			1868.7	3686.3		

†Refers to marginal effects on exit-specific probabilities (see previous section). Corresponding standard errors have been bootstrapped with 500 repetitions. ‡Includes 15 year dummies and 6 dummies for the previous sector of activity.
Standard errors for hazard ratios are robust with respect to clustering at the level of labour market regions.
Significance levels : † : 10% * : 5% ** : 1%

Table 2.3: Estimated hazard ratios and marginal effects for the local employment hazard (h_l) and the migration hazard (h_d) for the unstratified (UPLE) and the stratified partial likelihood estimator (SPLE), IABS 1983-1997, Women

	UPLE		**SPLE**			
	$exp(\hat{\beta}_l)$	$exp(\hat{\beta}_d)$	$exp(\hat{\beta}_l)$	$exp(\hat{\beta}_d)$	$\hat{\kappa}_l{}^{\dagger}$ (%)	$\hat{\kappa}_d{}^{\dagger}$ (%)
$\Pi_k(730\|\bar{x})(\%)$					42.2	4.3
Individual characteristics						
Married	0.883**	0.802**	0.882**	0.801**	-7.3*	-15.3*
Children	0.991	0.602**	0.991	0.623**	1.6	-36.0*
Age 26-29	0.816**	1.381**	0.822**	1.366**	-13.7*	44.1*
Age 30-34	0.840**	1.055	0.846**	1.049	-11.1*	11.0*
Age 40-44	1.068*	0.963	1.074*	0.971	5.0*	-5.4
Age 45-49	0.999	0.877	0.999	0.874	0.4	-12.2*
Age 50-54	0.800**	0.559**	0.809**	0.559**	-12.4*	-38.9*
Age > 55	0.400**	0.184**	0.406**	0.180**	-46.8*	-75.6*
High school degree	0.873**	0.583**	0.878**	0.584**	-7.4*	-38.0*
Higher education	0.834**	1.076	0.842**	1.091	-12.9*	13.4*
Unskilled blue-collar	0.981	0.971	0.987	0.974	-0.7	-1.9
White-collar	0.846	1.178**	0.867**	1.190**	-9.8*	24.2*
1st & 2nd wage quintile	0.883**	0.708**	0.885**	0.704**	-6.3*	-25.0*
4th & 5th wage quintile	0.911**	1.374**	0.912**	1.370**	-8.1*	38.0*
Prev. job tenure	0.981**	0.980**	0.981**	0.980**	-1.0*	-1.0*
Previously unemployed	1.379**	1.025	1.348**	1.043	22.2*	-6.5
Total prev. unemp. dur.	0.988**	0.976*	0.989**	0.977*	-0.6*	-1.8*
Recall from prev. employer	1.663**	0.680**	1.597**	0.660**	35.2*	-45.1*
Max. UB 0 months	0.736**	1.164	0.729**	1.087	-20.5*	19.6
Max. UB 1-12 months	1.033	1.142*	1.029	1.145*	1.4	12.6*
Max. UB 13-24 months	1.050†	0.915	1.046†	0.917	3.1*	-9.7

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... table 2.3 continued

Max. UB > 24 months	0.809**	0.704*	0.788**	0.712*	-14.3*	-21.9*
Income replacement rate	0.965**	0.990	0.965**	0.983	-0.6*	1.4*
Fourth quarter	0.773**	0.857**	0.774**	0.842**	-15.8*	-7.5*
Year and sector dummies‡	X	X	X	X		
Regional covariates						
UV-ratio	0.923**	0.927	0.949*	0.936	-3.3*	-4.4
Relative uv-ratio	0.936**	1.009	0.926**	1.013	-5.2*	4.1
Wage level	1.040	1.224**	1.025	1.225**	1.0	20.5*
Relative wage level	0.967*	0.901**	0.967*	0.910**	-2.0*	-7.6*
Employment growth	1.032**	0.979	1.037**	0.972	2.5*	-4.0
Accommodation ratio	0.975	0.903**	0.982	0.883**	-0.8	-10.7*
Male unemployment	1.025	0.945†	1.004	0.971	0.4	-3.0
Share of HS emp.	0.934**	0.993	1.001	0.974	0.2	-2.6
Employment level	1.009	1.017	0.899	1.407†	-8.8*	43.0*
Population density	0.965*	1.011	1.251**	1.008	15.7*	-7.6
Agriculture	0.997	0.951	1.006	0.861	1.0	-13.5*
Investment goods ind.	1.036	1.095	0.928	0.775	-3.9	-19.4
Consumption goods ind.	1.052†	1.104†	0.988	1.015	-0.8	2.0
Construction	1.026	1.074†	0.970	1.045	-1.9	5.7
Services	1.088*	1.195**	1.012	0.956	1.2	-4.5
Number of spells	50,585	50,585	50,585	50,585		
Number of exits	25,599	2,908	25,599	2,908		
Log-likelihood	-254,706.9	-27,826.5	-131,916.8	-14,440.1		
$\chi^2(df)$ clustering test			1430.8	122.2		

†Refers to marginal effects on exit-specific probabilities (see previous section). Corresponding standard errors have been bootstrapped with 500 repetitions. ‡Includes 15 year dummies and 6 dummies for the previous sector of activity.
Standard errors for hazard ratios are robust with respect to clustering at the level of labour market regions.
Significance levels : † : 10% * : 5% ** : 1%

Regional covariates The theoretical framework predicts that unemployed jobseekers choose search strategies that favour leaving regions with unfavourable re-employment opportunities or lower wage levels compared with other regions. Moreover, the relative attractiveness of the local area with respect to its living conditions may also affect job search strategies. Since amenities are mostly unobserved, estimation results for the unstratified estimator (UPLE) that is subject to unobserved regional heterogeneity could be biased if amenities compensate for unfavourable labour market conditions. Consistent with the idea that relatively unfavourable labour market conditions may be compensated for by relatively high regional amenities, the estimated effect of the relative unemployment-vacancy ratio on the migration hazard seems to be downward biased when omitting unobserved regional heterogeneity. While UPLE finds no evidence that men have higher migration hazards in regions with a relatively poor labour demand, the stratified estimator (SPLE) that mitigates biases from omitting unobserved regional heterogeneity suggests that men experience both a significantly lower local employment hazard and a higher migration hazard. As a consequence, an increase in the relative uv-ratio by one standard deviation significantly increases the cumulative migration probability by 12.2% and reduces the cumulative probability of exiting locally by 2.4%. By contrast, women experience only a reduced local employment hazard as a response to a higher relative uv-ratio. The missing positive effect on the migration hazard suggests that the shift of search effort to the distant region is too weak to counterbalance the negative effect of higher reservation wages that is induced by a higher relative uv-ratio.[20] This would be in line with the notion that women often are mobility constrained due to being a tied mover whose mobility decisions strongly depend on the situation of the male breadwinner. Compared to men, women thus are more dependent on local labour market conditions. In a region with a one standard deviation increase of the relative uv-ratio, the cumulative probability of leaving unemployment to any kind of employment within two years of job search decreases by around $2pp$ for women while the corresponding decrease is only around $1pp$ for men. For men, a higher migration probability partially compensates for a relatively weak local labour demand. Thus, at least for men, we can conclude that contrary

[20] As discussed in section 2.3, a higher relative uv-ratio reflects improving labour demand conditions in the distant region while keeping the conditions in the local region constant. This may raise reservation wages in all labour markets and may shift search effort to the distant labour market. The effect on the migration hazard is thus indeterminate as a negative reservation wage effect contrasts a positive direct effect and a positive search effort effect.

to the findings by Windzio (2004) regarding mobility between north and south Germany and in line with findings by Decressin (1994) on interregional migration flows in western Germany, unemployed individuals in western Germany choose search strategies that favour leaving regions with relatively weak labour demand. These results contradict the competing risk studies by Kettunen (2002) and Yankow (2002) who find no evidence that unemployed jobseekers react to local job-finding conditions.[21] The comparison between the UPLE and the SPLE estimates suggest, however, that their results may be subject to biases due to unobserved regional heterogeneity.

In contrast to the relative uv-ratio, the effect of the relative wage level is quite robust across both model specifications. This may be because central wage bargaining in Germany leaves little scope for local wage agreements. As a consequence, unemployment rather than wage differentials might compensate for differences in living conditions in Germany and unobserved regional heterogeneity may not have much of an effect. In any case, both model specifications indicate that the migration hazard is significantly lower in regions with a high relative wage level so that the cumulative probability of migration decreases by 6.6% (7.6%) for men (women). The local exit hazard, on the other hand, is either unaffected in the case of men or, somewhat surprisingly, marginally decreases in the case of women. This contradicts the idea that jobseekers shift search effort to the local area as a response to an increasing relative wage level. The reduced effect on the migration hazard is thus likely to mainly reflect the direct negative effect of lower wages in the distant labour market. On the other hand, the findings may indicate that the wage measures included in the analysis do not capture the behaviourally relevant interregional real wage differential. In addition, regions with relatively high wages may attract additional job searchers from other regions. In this case, an additional congestion effect may yield insignificant or negative effects on the local exit hazard even if there was an increase in local search effort.

Exit hazards may also be affected by changing aggregate conditions between 1983 and 1997. As discussed in section 2.3, a higher local unemployment-vacancy ratio when controlling for the relative uv-ratio indicates deteriorating overall labour demand conditions. The corresponding negative effect on the exit hazards seems to dominate a counteracting de-

[21] Yankow (2002) uses the unemployment rate, while Kettunen (2002) uses the unemployment-vacancy ratio as an indicator of local job-finding conditions.

crease in reservation wages and thus results in lower employment hazards everywhere in the economy which prolongs unemployment. Similarly, a higher aggregate wage level as reflected by a higher local wage level when controlling for interregional wage differentials results in significantly higher migration hazards, but somewhat surprisingly leaves the local exit hazard either unaffected in the case of women or even results in a lower local exit hazard for men. This latter finding might indicate that in addition to an increasing reservation wage, there is an encouragement effect of boom periods that shifts search effort from the local to the non-local labour market.

Some of the other covariates that have been included to control for additional local labour market characteristics also significantly affect the search strategy of the jobseeker and thus deserve some discussion. Since differences between the UPLE and the SPLE estimates are rather marginal, I discuss these findings based on the SPLE results. Higher employment growth, for example, significantly increases the local employment hazard, while there is a negative but insignificant effect on the migration hazard for both men and women. Thus, in line with Yankow (2002), the evidence in favour of a significant change in the allocation of search effort across regions as a response to local employment growth is inconclusive.

Secondly, the local accommodation of work creation schemes does not seem to have any effect on the migration hazard for men, but significantly reduces the migration hazard for women by almost 12%. Although the effect of the local accommodation ratio on the local exit hazard cannot be interpreted causally because the effect may be endogenous (see section 2.3), the significant effect on the female migration hazard can only be explained by a shift of search effort away from non-local employment. Thus, complementing the evidence on a regional locking-in effect of participating in work creation schemes (Lindgren and Westerlund, 2003), this paper presents evidence that such programs also have a locking-in effect on female non-participants by reducing search effort in non-local labour markets. The fact that such locking-in effects are only detectable for women may imply that participating in such programs is a more attractive substitute for regular non-local employment for women than for men as women tend to be more dependent on local labour market conditions.

As expected, a high share of male unemployment seems to reflect additional congestion effects on the labour market for men because the local employment hazard for men is significantly reduced while no effect can be found for women. Interestingly, there is also a

significant and strong negative effect on the male migration hazard such that the marginal effect on the cumulative migration probability even amounts to -18.2%. As an explanation, massive de-industrialisation and the corresponding increase in male unemployment may affect individual labour market behaviour by reducing the willingness to make efforts to regain employment if peers are also unemployed (Ritchie et al., 2005). Thus, both local and non-local exit hazards may be particularly low in these regions, resulting in a growing share of long-term unemployed.

The remaining covariates mostly have a minor impact only. A high share of employment in certain sectors such as agriculture, construction and services positively affects local exit hazards for men, a result that likely reflects the seasonal character of jobs in these sectors. No such effects can be found for women though. Moreover, a higher population density strongly accelerates local job exits among women, a result that, among others, may reflect that such regions offer more jobs that are easily accessible for women such as part-time employment.

2.4.2 Mobility effects by educational attainment

In order to examine whether skill groups differ in their responsiveness to labour market conditions, Table 2.4 compares estimates for low-skilled men with a high-school degree with estimates for skilled men who have either a vocational training or even a tertiary education.[22] I restrict the following discussion to men because low-skilled women have extremely low migration rates and the estimates do not therefore appear very reliable.

To begin with, it is important to note that the average skilled worker is four times more likely to leave the local labour market within two years of job search than the average low-skilled worker as can be seen from $\Pi_d(730|\bar{x})$. Moreover, skilled individuals are not only more mobile, but they are also more responsive to labour market conditions. In particular, skilled individuals have a significantly higher migration hazard in regions with a relatively weak labour demand and relatively low wages. The simultaneous reduction in the local employment hazard in response to a higher relative uv-ratio suggests that the increase in the migration hazard is partially due to a shift in search effort from the local to the distant region. By contrast, a similar though only weakly significant decrease of the local employment

[22]The share of unemployed people with tertiary education is around 5%.

hazard for low-skilled individuals in response to a higher relative uv-ratio is not reflected in a significantly increasing migration hazard. As a consequence, a one standard deviation increase of the relative uv-ratio results in a reduced probability of leaving unemployment to any type of employment by 1.5*pp* for low-skilled jobseekers, but only by 0.8*pp* in the case of skilled individuals. Low-skilled jobseekers thus constitute an immobile labour market segment that is more dependent on local labour market conditions than their skilled counterparts.

At first sight, labour market institutions such as passive and active labour market measures do not seem to explain the immobility of low-skilled as compared to skilled jobseekers. The length of UB entitlements shows similar result patterns for both skill groups with only extremely long entitlements having a somewhat stronger mobility-reducing impact on low-skilled individuals. Moreover, an extensive local supply of work creation programs increases the migration probability among low-skilled jobseekers by 30% (0.3*pp*). Thus, instead of a locking-in effect, low-skilled individuals seem to be pushed out of the labour market, a result that may indicate a crowding out of low-skilled jobseekers if work creation schemes tend to substitute for low-skilled jobs, but are primarily taken by skilled individuals. Labour market institutions thus do not seem to provide an explanation for lower mobility levels among low-skilled individuals. However, there are a number of observations that indirectly point towards the impact of the unemployment compensation and welfare system. One major drawback of the IAB-REG is that we cannot identify individuals who receive social benefits in addition to unemployment compensation and thus reach actual income replacement rates that are close to 100%[23]. This is because the IAB-REG does not include sufficient information on the household context (e.g. spouse income, number of dependent children, private savings) to identify individuals who are eligible for additional social benefits. A likely disincentive effect stemming from an unobserved receipt of supplementary welfare should thus be partially reflected in those individual-level characteristics that are related to the receipt of welfare such as low pre-unemployment wage income or having children. In particular, having dependent children strongly increases the likelihood of receiving social benefits, especially among low-skilled with rather low pre-unemployment earnings. The relatively weak effect

[23]Supplementary social benefits may thus explain why the estimated effect for the income replacement rate that refers exclusively to the unemployment compensation in Table 2.4 shows only weak effects for low-skilled individuals who are more likely to be unemployed social benefit recipients.

of having children compared to the strong mobility-enhancing impact on skilled individuals may thus partially reflect these disincentive effects. Similarly, pre-unemployment earnings in the lowest two wage quintiles have a strong mobility-reducing impact, especially among low-skilled individuals who mainly (62.5%) earn a wage income that falls into these lower wage quintiles. The findings thus provide some indirect evidence that the unemployment compensation and welfare system may provide at least a partial explanation why low-skilled jobseekers in Germany are less mobile than their skilled counterparts.

Table 2.4: Estimated hazard ratios and marginal effects for the local employment hazard (h_l) and the migration hazard (h_d) for a stratified partial likelihood estimator by skill-level, IABS 1983-1997, Men

	Low-skilled				Skilled			
	$exp(\hat{\beta}_l)$	$exp(\hat{\beta}_d)$	$\hat{\kappa}_l$ (%)	$\hat{\kappa}_d$ (%)	$exp(\hat{\beta}_l)$	$exp(\hat{\beta}_d)$	$\hat{\kappa}_l{}^{\dagger}$ (%)	$\hat{\kappa}_d{}^{\dagger}$ (%)
$\Pi_k(730\|\bar{x})$			50.7	1.1			59.3	4.0
Individual characteristics								
Married	1.209**	1.249	11.8 *	16.7*	1.214**	1.219**	9.8*	11.5*
Children	1.029	1.073	1.6	6.1*	1.142**	1.407**	5.3*	30.9*
Age 26-29	1.025	1.004	1.6	-0.5	1.113**	1.101*	5.3*	4.8
Age 30-34	1.006	0.794	1.0	-20.3*	1.053**	1.101*	2.1*	7.2*
Age 40-44	0.927	0.959	-4.3	-1.5	0.966	0.921†	-1.4	-6.2*
Age 45-49	0.892*	0.755†	-6.0 *	-21.0*	0.920**	0.866*	-3.7*	-9.9*
Age 50-54	0.817**	0.453**	-10.6 *	-51.0*	0.883**	0.738**	-5.4*	-21.6*
Age > 55	0.482**	0.190**	-36.1 *	-75.8*	0.381**	0.211**	-41.0*	-69.0*
Tertiary education		n/a			0.763**	1.254**	-20.3*	30.4*
Unskilled blue-collar	0.926*	0.843	-4.0 *	-13.1*	0.874**	0.888**	-6.3*	-5.7*
White-collar	0.762**	1.462*	-18.4 *	55.1*	0.609**	1.235**	-27.3*	44.5*
1st & 2nd wage quintile	0.787**	0.648**	-12.4 *	-28.9*	0.826**	0.855**	-8.6*	-6.4*
4th & 5th wage quintile	1.018	1.033	0.9	2.5	0.981	1.403**	-2.8*	39.2*
Prev. job tenure (in quarters)	0.978**	0.972**	-1.0 *	-1.7*	0.985**	0.964**	-0.4*	-2.7*
Previously unemployed	1.270**	0.901	16.5 *	-16.4*	1.257**	0.964	14.5*	-11.2*
Total prev. unemp. dur. (in qu.)	0.980**	0.949**	-1.1 *	-4.3*	0.985**	0.967**	-0.6*	-2.5*
Recall from prev. employer	1.548**	0.598**	29.1 *	-48.9*	1.408**	0.545**	20.8*	-52.4*
Max. UB 0 months	0.634**	0.783	-25.7 *	-10.2*	0.535**	0.681**	-31.0*	-14.9*
Max. UB 1-12 months	0.827**	0.908	-11.0 *	-3.3	0.724**	1.100*	-17.4*	23.5*
Max. UB 13-24 months	0.981	0.891	-1.0	-10.2	1.005	0.975	0.3	-2.7
Max. UB > 24 months	0.870†	0.430*	-7.0 *	-54.5*	0.882**	0.675**	-4.7*	-27.7*

Continued on next page...

... table 2.4 continued

Income replacement rate	0.983†	0.968	-0.3	-2.0	0.971**	0.939**	0.3*	-3.2*
Fourth quarter	0.465**	0.719**	-39.1 *	-9.3	0.524**	0.746**	-30.5*	-4.3
Year and sector dummies‡	X	X			X	X		
Regional covariates								
UV-ratio	0.897**	0.959	-6.4*	-0.5	0.880**	0.881**	-6.0*	-6.6*
Relative uv-ratio	0.950†	1.067	-3.2*	8.4	0.970*	1.120**	-2.3*	12.7*
Wage level	0.945	1.069	-3.6*	8.7	0.950**	1.050†	-2.9*	7.0*
Relative wage level	0.975	0.947	-1.4	-4.4	0.994	0.929**	0.1	-6.6*
Employment growth	1.055**	0.954	3.3*	-6.4	1.041**	0.972	2.2*	-4.3
Accommodation ratio	1.017	1.313**	0.5	30.0*	1.055**	1.009	2.9*	-1.3
Male unemployment	0.921**	0.919	-4.9*	-5.5	0.913**	0.776**	-4.1*	-8.9*
Share of HS emp.	1.013	0.599**	2.0	-39.8*	1.029	0.972	1.7	-3.8
Employment level	1.202	1.526	9.9	41.2	1.114	0.978	5.6	-6.7
Population density	1.250	1.175	13.2	7.8	0.901	0.954	-5.7	-0.8
Agriculture	1.041	0.892	2.7	-11.8	1.075**	1.182*	3.1*	13.9*
Investment goods ind.	1.008	1.625	-0.4	60.8	1.107†	1.120	5.0	6.9
Consumption goods ind.	0.958	1.003	-2.9	1.5	0.954	0.841	-2.0	-13.9*
Construction	1.120*	1.275	6.5*	22.1	1.092**	1.101	4.3*	5.7
Services	0.987	1.140	-1.1	14.2	1.185**	1.059	8.7*	-1.9
Number of spells	16,562	16,562			70,698	70,698		
Number of exits	10,444	667			46,508	6,115		
Log-likelihood	-40,633.6	-2,321.9			-248,563.0	-31,684.9		
$\chi^2(df)$ clustering test	474.7 (56)	80.3 (56)			1,073.1 (56)	186.2 (56)		

†Refers to marginal effects on exit-specific probabilities (see previous section). Corresponding standard errors have been bootstrapped with 500 repetitions. ‡Includes 15 year dummies and 6 dummies for the previous sector of activity.

Standard errors for hazard ratios are robust with respect to clustering at the level of labour market regions.

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

2.5 Conclusion

This paper has looked at a competing risks model of exiting unemployment to either a local or a non-local job in order to test whether unemployed individuals in western Germany adjust their search strategy to favour migration away from depressed regions. The equilibrating role of interregional migration critically hinges on such search strategies. Previous studies which have looked at the migratory behaviour of unemployed jobseekers in a competing risks framework have not found any evidence that individuals adjust their search strategies to local job-finding conditions (Kettunen, 2002; Yankow, 2002). The results of this paper indicate that this may be due to biases that result from unobserved regional heterogeneity. Using a stratified partial likelihood estimator that takes into account location-specific fixed effects and thus mitigates biases from unobserved regional heterogeneity, the paper comes to the following conclusions:

- Unemployed jobseekers are at least partially responsive to local labour market conditions. In particular, unemployed men choose search strategies that favour migration away from regions with a relatively weak labour demand compared to other regions. This latter effect is not detectable for a specification that is subject to unobserved regional heterogeneity. The implied downward bias is consistent with the idea that unobserved regional amenities compensate for unfavourable labour market conditions.

- Women are less responsive to labour market conditions than their male counterparts. Moreover, family-responsibilities strongly reduce migration among women, but not among men. These gender differences probably reflect that women are often tied movers and have a weaker labour force attachment than men. This may also explain why there is significant evidence that extensive local work creation schemes exercise a regional locking-in effect on women, but not on men. If women are more dependent on the local labour market, participation in active labour market programs may be a more attractive substitute for non-local employment for women than for men.

- Compared with skilled men, low-skilled men are relatively immobile and only weakly respond to labour market conditions. Deteriorating labour demand conditions thus translate into prolonged unemployment. These findings are unsettling if labour mobil-

ity is envisaged as one way of reducing regional disparities and of increasing employment levels.

- Even though labour market conditions are relevant for the search strategy of some groups of unemployed jobseekers, geographic mobility is clearly driven by individual characteristics such as age, educational attainment and employment history (e.g. previous recalls).

- The mobility-reducing effects of being entitled to a long receipt of unemployment benefits and generous income replacement rates suggest that the institutional framework is also a major determinant of the migratory behaviour of unemployed jobseekers. Such disincentives may also partially explain the weak responsiveness to regional labour market conditions among low-skilled individuals because this labour segment is more likely to receive social benefits in addition to unemployment compensation and thus reach extensive income replacement rates.

The responsiveness of skilled men to labour market conditions suggests that labour mobility in western Germany may at least contribute to the reduction of regional disparities although the low responsiveness among low-skilled individuals indicates that such counteracting forces are partially weak. Moreover, the equilibrating role of labour mobility ultimately depends not only on the responsiveness, but also on the level and speed of mobility. Even among the most mobile segments of the German labour force, mobility is rather low in an international comparison. The findings from this paper are thus in line with previous results from a study by Möller (1995) that adjustment processes after region-specific shocks tend to be slow in Germany. The results presented in this paper already point towards the unemployment compensation system as one possible reason for low mobility levels and a weak responsiveness to regional labour market conditions, but further research is necessary on this topic.

Appendix

A – Comparative statics for reservation wages

A.1 – Transfer payments

$$\frac{d\, w_l^r}{d\, b} = \frac{r}{\alpha_l \sigma(e_l)(1 - F_l(w_l^r)) + \alpha_d \sigma(e_d)(1 - F_d(w_d^r)) + r} > 0$$

and $\frac{d\, w_d^r}{d\, b} = \frac{d\, w_l^r}{d\, b}$ because $w_d^r = w_l^r + (a_l - a_d) + pc_d + r\ mc_d$ (see equation (2.6)).

A.2 – Job offer arrival rate

$$\frac{d\, w_l^r}{d\, \alpha_k} = \frac{\sigma(e_k)\ P(w \geq w_k^r)\ E(w|w \geq w_k^r)}{\alpha_l \sigma(e_l)(1 - F_l(w_l^r)) + \alpha_d \sigma(e_d)(1 - F_d(w_d^r)) + r} > 0 \qquad \forall k = d, l$$

and again, $\frac{d\, w_d^r}{d\, \alpha_k} = \frac{d\, w_l^r}{d\, \alpha_k}$.

A.3 – Search cost

$$\frac{d\, w_l^r}{d\, c_k(e_k)} = -\frac{r\ c_k'(e_k)}{\alpha_l \sigma(e_l)(1 - F_l(w_l^r)) + \alpha_d \sigma(e_d)(1 - F_d(w_d^r)) + r} < 0 \qquad \forall k = d, l$$

since $c_k'(e_k) > 0$. Again, $\frac{d\, w_d^r}{d\, c_k(e_k)} = \frac{d\, w_l^r}{d\, c_k(e_k)}$.

A.4 – Moving cost For flow-type moving cost, the effect on the local reservation wage is:

$$\frac{d\, w_l^r}{d\, pc_d} = -\frac{\alpha_d \sigma(e_d)(1 - F_d(w_d^r))}{\alpha_l \sigma(e_l)(1 - F_l(w_l^r)) + \alpha_d \sigma(e_d)(1 - F_d(w_d^r)) + r} < 0$$

and it follows that $-1 < \frac{d\, w_l^r}{d\, pc_d} < 0$. The effect on w_d^r can be derived from equation 2.6:

$$\frac{d\, w_d^r}{d\, pc_d} = \frac{d\, w_d^r}{d\, w_l^r}\frac{d\, w_l^r}{d\, pc_d} + 1 = 1 - \frac{\alpha_d(e_d)(1 - F_d(w_d^r))}{\alpha_l(e_l)(1 - F_l(w_l^r)) + \alpha_d(e_d)(1 - F_d(w_d^r)) + r} > 0$$

and it follows that $0 < \frac{d\, w_d^r}{d\, pc_d} < 1$. For lump-sum moving cost, the corresponding effects are $\frac{d\, w_l^r}{d\, mc_d} = r\frac{d\, w_l^r}{d\, pc_d} < 0$ and $\frac{d\, w_d^r}{d\, mc_d} = r\frac{d\, w_l^r}{d\, pc_d} > 0$.

A.5 – Amenities

$$\frac{d\, w_l^r}{d\, a_l} = -\frac{\alpha_d \sigma(e_d)(1 - F_d(w_d^r))}{\alpha_l \sigma(e_l)(1 - F_l(w_l^r)) + \alpha_d \sigma(e_d)(1 - F_d(w_d^r)) + r} < 0$$

$$\frac{d\, w_l^r}{d\, a_d} = \frac{\alpha_d \sigma(e_d)(1 - F_d(w_d^r))}{\alpha_l \sigma(e_l)(1 - F_l(w_l^r)) + \alpha_d \sigma(e_d)(1 - F_d(w_d^r)) + r} > 0$$

while $\frac{d\, w_d^r}{d\, a_l} = 1 + \frac{d\, w_l^r}{d\, a_l} > 0$ and $\frac{d\, w_d^r}{d\, a_d} = \frac{d\, w_l^r}{d\, a_d} - 1 < 0$ because $-1 < \frac{d\, w_l^r}{d\, a_l} < 0$ and $0 < \frac{d\, w_l^r}{d\, a_d} < 1$.

A.6 – Wage distribution Consider the following transformation of equation (2.6):

$$\begin{aligned}(r+\alpha_l\sigma(e_l)+\alpha_d\sigma(e_d))w_l^r &= r(b_l-c_l(e_l)-c_d(e_d)) \\ &\quad +\alpha_l\sigma(e_l)[E_{F_l}(w)+\int_0^{w_l^r}F_l(w)dw] \\ &\quad +\alpha_d\sigma(e_d)[E_{F_d}(w)+\int_0^{w_d^r}F_d(w)dw]\end{aligned} \tag{2.9}$$

For an alternative wage offer function G_l whose mean is μ units larger than the mean of F_l, i.e. G_l first order stochastically dominates F_l with $G_l(w+\mu)=F_l(w)$, equation (2.9) for the wage offer distribution G_l can be written as

$$\begin{aligned}(r+\alpha_l\sigma(e_l)+\alpha_d\sigma(e_d))w_l^r(\mu) &= r(b_l-c_l(e_l)-c_d(e_d)) \\ &\quad +\alpha_l\sigma(e_l)[E_{F_l}(w)+\mu+\int_0^{w_l^r(\mu)}F_l(w-\mu)dw] \\ &\quad +\alpha_d\sigma(e_d)[E_{F_d}(w)+\int_0^{w_d^r(\mu)}F_d(w)dw]\end{aligned} \tag{2.10}$$

$\lim_{\mu\to 0}$ (2.10)−(2.9) and rearranging yields:

$$\frac{d\,w_l^r}{d\,\mu}=\frac{\alpha_l\sigma(e_l)(1-F_l(w_l^r))}{\alpha_l\sigma(e_l)(1-F_l(w_l^r))+\alpha_d\sigma(e_d)(1-F_d(w_d^r))+r}>0$$

and $\frac{d\,w_l^r}{d\,\mu}=\frac{d\,w_d^r}{d\,\mu}$. For a wage offer distribution $G_d(w+\mu)=F_d(w)$, the analogous result is $\frac{d\,w_l^r}{d\,\mu}=\frac{d\,w_d^r}{d\,\mu}>0$.

B – Comparative statics for the allocation of search effort

The optimal allocation of search effort across regions e_k^* is derived by differentiating V_l^u with respect to e_k:

$$c_k'(e_k^*)=\frac{\alpha_k\sigma'(e_k^*)}{r}\int_{w_k^r}^{w_{max}}(w-w_k^r)dF_k(w)\qquad\forall k=d,l \tag{2.11}$$

The effect of changing conditions on the allocation of search effort is given as $\frac{d\,e_k^*}{d\,x}=-\frac{F_x}{F_{e_k^*}}$ with x being the changing condition of interest. It holds that

$$F_{e_k^*}=\alpha_k\sigma''(e_k^*)\int_{w_k^r}^{w_{max}}(w-w_k^r)dF_k(w)-c_k''(e_k^*)<0\qquad\forall k=d,l$$

because $\sigma''(e_k^*)<0$ and $c_k''(e_k^*)>0$. Thus, $sgn(\frac{d\,e_k^*}{d\,x})=sgn(F_x)$.

B.1 – Transfer payments

$$sgn[\frac{d\ e_k^*}{d\ b_l}] = sgn[-\frac{\alpha_k \sigma'(e_k^*)}{r}\ \ \frac{d\ w_k}{d\ b_l}\ \ (1 - F_l(w_k^r))] < 0 \quad \forall k = d, l$$

because $\sigma'(e_k^*) > 0$ and $\frac{d\ w_k}{d\ b_l} > 0$ follows from A.1.

B.2 – Job offer arrival rate

$$sgn[\frac{d\ e_l^*}{d\ \alpha_l}] = sgn[\frac{\sigma'(e_l^*)}{r} \int_{w_l^r}^{w_{max}} (w - w_l^r) dF_l(w)$$
$$[1 - \frac{\alpha_l \sigma(e_l^*)(1 - F_l(w_l^r))}{\alpha_l \sigma(e_l^*)(1 - F_l(w_l^r)) + \alpha_d \sigma(e_d^*)(1 - F_d(w_d^r)) + r}]] > 0$$

$$sgn[\frac{d\ e_l^*}{d\ \alpha_d}] = sgn[-\frac{\alpha_l \sigma'(e_l^*)}{r} \frac{d\ w_l^r}{d\ \alpha_d}(1 - F_l(w_l^r))] < 0$$

because $\frac{d\ w_l^r}{\alpha_d} > 0$ follows from A.2. Analogously, it can be shown that $\frac{d\ e_d^*}{d\ \alpha_d} > 0$ and $\frac{d\ e_d^*}{d\ \alpha_l} < 0$.

B.3 – Search cost

$$sgn[\frac{d\ e_l^*}{d\ c_l}] = sgn[-\frac{\alpha_l \sigma'(e_l^*)}{r} \frac{d\ w_l^r}{d\ c_l}(1 - F_l(w_l^r)) - c_l''(e_l^*)]$$

which is ambiguous because $c_k(e_k^*)$ is convex with $c_k''(e_k^*) > 0$ and $\frac{d\ w_l^r}{d\ c_l} < 0$. Analogously, $\frac{d\ e_d^*}{d\ c_d}$ is also ambiguous. By contrast, increasing search costs in a region increases search effort in alternative regions, i.e.

$$sgn[\frac{d\ e_l^*}{d\ c_d}] = sgn[-\frac{\alpha_l \sigma'(e_l^*)}{r} \frac{d\ w_l^r}{d\ c_d}(1 - F_l(w_l^r))] > 0$$

because $\frac{d\ w_l^r}{d\ c_d} < 0$ follows from A.3. For the same reason, it can be shown that $\frac{d\ e_d^*}{d\ c_l} > 0$.

B.4 – Moving cost

$$sgn[\frac{d\ e_l^*}{d\ pc_d}] = sgn[-\frac{\alpha_l \sigma'(e_l^*)}{r} \frac{d\ w_l^r}{d\ pc_d}(1 - F_l(w_l^r))] > 0$$

$$sgn[\frac{d\ e_d^*}{d\ pc_d}] = sgn[-\frac{\alpha_d \sigma'(e_d^*)}{r} \frac{d\ w_d^r}{d\ pc_d}(1 - F_d(w_d^r))] < 0$$

because $\frac{d\ w_l^r}{d\ pc_d} < 0$ and $\frac{d\ w_d^r}{d\ pc_d} > 0$ follows from A.4. Lump-sum moving costs affect the allocation of search effort analogously with $\frac{d\ e_l^*}{d\ mc_d} > 0$ and $\frac{d\ e_d^*}{d\ mc_d} < 0$.

B.5 – Amenities

$$sgn[\frac{d\ e_l^*}{d\ a_l}] = sgn[-\frac{\alpha_l \sigma'(e_l^*)}{r}\frac{d\ w_l^r}{d\ a_l}(1 - F_l(w_l^r))] > 0$$

$$sgn[\frac{d\ e_l^*}{d\ a_d}] = sgn[-\frac{\alpha_l \sigma'(e_l^*)}{r}\frac{d\ w_l^r}{d\ a_d}(1 - F_l(w_l^r))] < 0$$

because $\frac{d\ w_l^r}{d\ a_l} < 0$ and $\frac{d\ w_l^r}{d\ a_d} > 0$ follows from A.5. It also follows from A.5 that $\frac{d\ e_d^*}{d\ a_d} > 0$ and $\frac{d\ e_d^*}{d\ a_l} < 0$.

B.6 – Wage distribution As in A.6, consider an improvement of the local relative to the distant wage offer distribution, i.e. $G_l(w + \mu) = F_l$. After transforming equation (2.11) as discussed in A.6 and taking the difference between the resulting equation and the equation for the original wage distribution F_l, $\lim_{\mu \to 0}$ yields:

$$sgn[\frac{d\ e_l^*}{d\ \mu}] = sgn(\frac{\alpha_l \sigma'(e_l^*)}{r}[(1 - F_l(w_l^r))(1 - \frac{d\ w_l^r}{d\ \mu})]) > 0$$

because $0 < \frac{d\ w_l^r}{d\ \mu} < 1$. By contrast, search effort in the distant region decreases because

$$sgn[\frac{d\ e_d^*}{d\ \mu}] = sgn[-\frac{\alpha_d \sigma'(e_d^*)}{r}[(1 - F_d(w_l^r))\frac{d\ w_d^r}{d\ \mu}]] < 0$$

follows from $\frac{d\ w_d^r}{d\ \mu} > 0$. The analogous opposite effects can be shown in case of an improvement in the distant relative to the local wage distribution.

C – Data Description

C.1 - Description and data sources of regional covariates

Variables	Description[a]	Data Source[b]	1yr lag
Regional covariates with area variation			
UV-ratio	Ratio between registered unemployed and vacancies, i.e. U_l/V_l	BA	X
Relative uv-ratio	Local uv-ratio divided by uv-ratio in all distant regions, i.e. $\frac{U_l/V_l}{U_d/V_d}$	BA	X
Accommodation ratio	Number of participants in work creation schemes (*Arbeitsbeschaffungsmaßnahmen*) divided by number of unemployed people plus participants in such schemes	BA	X
Share of sector j	Employment share in agriculture, investment goods industry, consumption goods industry, construction, services (reference: retail)	IABS-REG	X
Share of HS employment	Share of employees with a tertiary education	IABS-REG	X
Male unemployment	Share of unemployed people who are male	BA	X
Employment level	Total average employment	IABS-REG	X
Population density	Population divided by area	New Chronos, BA	X
Regional covariates with area and individual variation			
Wage level	Individual *i's* predicted wage level in the local labour market region (see C.3 for details)	IAB-R01	
Relative wage level	Individual *i's* predicted wage level in the local labour market divided by individual *i's* average aggregate wage level (see C.3 for details)	IAB-R01	
Employment growth	Average employment change compared with the previous year in individual *i's* skill group (low-skilled without vocational training, skilled with vocational training or tertiary education)	IABS-REG	

[a] All regional indicators have been aggregated to the level of labour market regions.
[b] See C.2 for details on the data sources.

C.2 - Details on regional data sources

- BA - Data is released by the Federal Employment Agency (*Bundesagentur für Arbeit*) and is coded at the level of labour office agencies (*Arbeitsamtsbezirke*). Since the IABS-REG only contains information on the microcensus region of the workplace and microcensus regions and labour office agencies are spatially misaligned, all BA data have been interpolated to microcensus regions by using the simple area weighting that has been proposed in chapter (1) of this dissertation.

- New Chronos - Regional database that is released by Eurostat.

- IABS-REG - Some regional covariates have been calculated based on the IABS-REG itself. Employment in region k by sector and skill-group is computed for the 15th of each month. Employment shares by sector and high-skilled individuals are then calculated as a yearly average across all months. Employment growth by skill-group refers to the change of these yearly averages.

- IAB-R01 - The IAB-R01 is released by the IAB (*Institut für Arbeitsmarkt- und Berufsforschung*) and has the same spell structure as the IABS-REG which is described in detail in section 2.3. In contrast to the IABS-REG, the sample contains 2.2% instead of 1% of the population working in a job that is subject to social insurance contributions and contains fewer individual-level information.

C.3 - Wage estimates based on the IAB-R01 For the wage estimates, I create yearly sub-samples with all full-time employees on 01/01. These sub-samples contain between 200-300,000 observations for each year with a minimum of 300 observations in each region. For each year, I separately estimate Mincerian type wage equations with years of education, experience, squared experience, nine sector dummies and five occupation type dummies. I take account of the fact that wages in the IAB-R01 are top-coded and run a censored regression instead of simple OLS. Based on the resulting estimates, I predict the average aggregate wage in year t for each individual i in the IAB-REG sample (IAB-REG and IAB-R01 are comparable in the used covariates). I run the same type of wage regression separately for each labour market region and year to predict individual i's wage in year t in the local region. Dividing this measure by the predicted average aggregate wage for individual i in year t yields the relative wage measure.

C.4 - Descriptive statistics - Men

			Spells ending ...			
	All Spells		in migration		locally	
Variables	Mean	SD	Mean	SD	Mean	SD
Individual level covariates						
Family background[1]						
Married	0.62	0.48	0.58	0.49	0.62	0.49
Children	0.56	0.50	0.53	0.50	0.56	0.50
Age at beginning of unemployment spell[1] (Reference: 35-39 years)						
26-29	0.20	0.40	0.23	0.42	0.23	0.42
30-34	0.19	0.39	0.25	0.43	0.20	0.40
40-44	0.11	0.32	0.13	0.34	0.12	0.33
45-49	0.12	0.32	0.11	0.32	0.13	0.33
50-54	0.10	0.31	0.07	0.26	0.11	0.31
> 55	0.14	0.34	0.03	0.17	0.07	0.26
Formal education[1] (Reference: Vocational training)						
High school degree	0.19	0.39	0.10	0.30	0.18	0.39
Higher education	0.04	0.21	0.13	0.34	0.03	0.16
Previous wage quintile[1] (Reference: 3rd quintile)						
1st & 2nd	0.49	0.50	0.44	0.50	0.49	0.50
4th & 5th	0.32	0.47	0.39	0.49	0.30	0.46
Previous job status[1] (Reference: Skilled blue-collar worker)						
Unskilled blue-collar	0.37	0.48	0.29	0.45	0.38	0.48
White-collar	0.16	0.37	0.34	0.48	0.11	0.31
Previous sector of activity[1] (Reference: Retail)						
Agriculture	0.05	0.22	0.02	0.15	0.06	0.24
Investment goods ind.	0.23	0.42	0.19	0.39	0.19	0.40
Consumption goods ind.	0.08	0.28	0.08	0.27	0.08	0.26
Construction	0.27	0.44	0.20	0.40	0.33	0.47
Services	0.21	0.41	0.30	0.46	0.19	0.39
Max. entitlement period for unemployment benefits[1] (Reference: 12 months)						
0 months	0.06	0.24	0.11	0.31	0.06	0.23
1-12 months	0.14	0.35	0.19	0.39	0.15	0.35
13-24 months	0.14	0.35	0.12	0.32	0.15	0.36
> 24 months	0.16	0.36	0.05	0.23	0.10	0.30

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... table 2.5 continued

Other individual characteristics						
Income replacement rate (in %)	65.1	3.5	64.5	3.8	65.2	3.4
Job tenure prev. job (quarters)	10.2	16.5	6.2	9.7	6.8	10.9
Previously unemployed[1]	0.68	0.47	0.61	0.49	0.77	0.42
Total prev. unemp. dur. (quart.)	3.9	5.4	3.1	4.8	4.4	5.4
Recall from prev. employer[1]	0.27	0.45	0.11	0.31	0.35	0.48
Regional covariates with area variation[2]						
UV-ratio	0.20	1.2	0.01	1.0	0.13	1.0
Relative uv-ratio	0.13	1.1	0.08	1.1	0.07	1.0
Accommodation ratio	0.00	1.0	-0.02	1.0	0.03	1.0
Male unemployment	0.28	1.0	0.20	0.9	0.37	1.0
Share of HS employment	0.36	1.2	0.67	1.2	0.28	1.2
Employment level	1.00	1.9	1.31	2.1	0.82	1.9
Population density	0.30	1.2	0.39	1.1	0.18	1.1
Employment shares by sector (Reference: Retail)						
Agriculture	0.09	1.2	0.01	1.1	0.07	1.1
Investment goods ind.	-0.14	0.8	-0.18	0.8	-0.17	0.8
Consumption goods ind.	-0.23	1.1	-0.40	0.9	-0.12	1.1
Construction	-0.25	0.9	-0.33	0.8	-0.21	0.9
Services	0.19	1.0	0.38	0.9	0.19	1.0
Regional covariates with area and individual variation[2]						
Wage level	0.54	0.82	0.73	0.9	0.47	0.7
Relative wage level	0.10	0.82	0.11	0.8	0.09	0.8
Employment growth	0.24	0.75	0.26	0.7	0.29	0.8
Number of spells	87,260		6,782		56,952	

[1]Dummy Variable(s)

[2]Time-varying covariates with quarterly (uv-ratio, relative uv-ratio, accommodation ratio, male unemployment) or yearly (all other covariates) variation. All regional covariates are continuous and have been standardised at the regional level. Thus, a value of one means one standard deviation above the average labour market region between 1983 and 1997.

C.5 - Descriptive statistics - Women

			Spells ending ...			
	All Spells		in migration		locally	
Variables	Mean	SD	Mean	SD	Mean	SD
Individual level covariates						
Family background[1]						
Married	0.65	0.48	0.46	0.50	0.59	0.49
Children	0.53	0.50	0.37	0.48	0.53	0.50
Age at beginning of unemployment spell[1] (Reference: 35-39 years)						
26-29	0.23	0.42	0.33	0.47	0.22	0.41
30-34	0.20	0.40	0.24	0.43	0.20	0.40
40-44	0.12	0.33	0.12	0.32	0.15	0.35
45-49	0.11	0.31	0.09	0.28	0.13	0.33
50-54	0.10	0.30	0.06	0.23	0.10	0.30
> 55	0.10	0.31	0.02	0.14	0.05	0.22
Formal education[1] (Reference: Vocational training)						
High school degree	0.26	0.44	0.12	0.32	0.23	0.42
Higher education	0.05	0.22	0.13	0.33	0.05	0.21
Previous wage quintile[1] (Reference: 3rd quintile)						
1st & 2nd	0.83	0.37	0.69	0.46	0.84	0.37
4th & 5th	0.09	0.29	0.20	0.40	0.08	0.28
Previous job status[1] (Reference: Skilled blue-collar worker)						
Unskilled blue-collar	0.28	0.45	0.18	0.39	0.29	0.45
White-collar	0.40	0.49	0.59	0.49	0.38	0.49
Previous sector of activity[1] (Reference: Retail)						
Agriculture	0.02	0.13	0.01	0.09	0.02	0.15
Investment goods ind.	0.14	0.35	0.10	0.30	0.11	0.31
Consumption goods ind.	0.15	0.36	0.08	0.28	0.15	0.35
Construction	0.02	0.14	0.02	0.12	0.02	0.15
Services	0.44	0.50	0.55	0.50	0.47	0.50
Max. entitlement period for unemployment benefits[1] (Reference: 12 months)						
0 months	0.06	0.23	0.09	0.28	0.07	0.25
1-12 months	0.15	0.36	0.17	0.38	0.17	0.37
13-24 months	0.14	0.35	0.10	0.30	0.16	0.36
> 24 months	0.14	0.34	0.05	0.21	0.09	0.28

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... table 2.6 continued

Other individual characteristics						
Income replacement rate (in %)	65.0	3.4	63.9	3.5	64.8	3.5
Job tenure prev. job (quarters)	12.4	16.8	8.8	11.7	8.3	12.7
Previously unemployed[1]	0.55	0.50	0.54	0.50	0.65	0.48
Total prev. unemp. dur. (quart.)	2.5	4.0	2.32	4.0	3.0	4.2
Recall from prev. employer[1]	0.21	0.40	0.10	0.30	0.27	0.44
Regional covariates with area variation[2]						
UV-ratio	0.01	1.0	-0.14	0.9	-0.10	0.9
Relative uv-ratio	0.08	1.1	0.03	1.1	0.02	1.1
Accommodation ratio	-0.01	1.0	-0.08	0.9	-0.00	1.0
Male unemployment	0.08	0.9	0.13	0.9	0.13	1.0
Share of HS employment	0.51	1.2	0.84	1.2	0.53	1.2
Employment level	1.1	2.0	1.47	2.2	1.06	2.0
Population density	0.36	1.1	0.46	1.2	0.28	1.1
Employment shares by sector (Reference: Retail)						
Agriculture	0.01	1.1	-0.10	0.9	-0.04	1.0
Investment goods ind.	-0.11	0.9	-0.18	0.8	-0.17	0.8
Consumption goods ind.	-0.33	1.0	-0.47	0.9	-0.30	1.0
Construction	-0.34	0.8	-0.45	0.7	-0.33	0.8
Services	0.28	1.0	0.50	1.0	0.35	1.0
Regional covariates with area and individual variation[2]						
Wage level	-0.55	0.7	-0.44	0.7	-0.50	0.7
Relative wage level	-0.19	1.2	-0.19	1.2	-0.22	1.2
Employment growth	0.16	0.8	0.25	0.7	0.19	0.8
Number of spells	50,585		2,908		25,599	

[1]Dummy Variable(s)

[2]Time-varying covariates with quarterly (uv-ratio, relative uv-ratio, accommodation ratio, male unemployment) or yearly (all other covariates) variation. All regional covariates are continuous and have been standardised at the regional level. Thus, a value of one means one standard deviation above the average labour market region between 1983 and 1997.

Appendix E - Maximum entitlement periods for unemployment benefits by age

	up to age							
Period	42	44	45	47	49	52	54	57
until 01/1995	12	12	12	12	12	12	12	12
01/1985 - 12/1985	12	12	12	12	18	18	18	18
1986 - 06/1987	12	16	16	16	20	20	24	24
07/1987 - 03/1997	18	20	22	22	26	26	32	32

Source: Plaßmann (2002)

Chapter 3

Unemployment Duration in Germany: Individual and Regional Determinants of Local Job Finding, Migration and Subsidised Employment

joint with Ralf A. Wilke[1]

Abstract

Recent labour market reforms in Germany aim, among other things, at reducing unemployment by restricting passive unemployment measures, emphasising local labour market policies and re-structuring public employment services. This paper uses extensive individual administrative and regional aggregate data to explore the extent to which these factors are likely to contribute to the shortening of unemployment duration. For this purpose, we estimate a semi-parametric duration model with three competing exit states in order to disentangle the relevance of these factors for exits to regular local, regular non-local and subsidised employment. Our results suggest that changes in the unemployment compensation system rather than local employment policies and increasing job counselling efforts may accelerate exits to regular employment.

Keywords: competing-risk, labour market policy, individual and regional data
JEL: J64, J61, J68

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3.1 Introduction

Throughout the last two decades Germany has experienced persistently high and even rising levels of unemployment. At the same time, the share of unemployed who remain unemployed even after one year of job search has also gone up to 37% in 2002 according to administrative data from the Federal Employment Agency (Bundesagentur für Arbeit , 2002). Survey-based measures even report shares of long-term unemployment of close to 50% (Machin and Manning, 1999). This is much higher than in the US, but reflects a labour market situation that is not uncommon to many European countries. In this context, improved knowledge of how individual characteristics as well as the regional and institutional context shape labour market outcomes of unemployed jobseekers is of central concern to policy makers aiming to design policies that will contribute to a shortening of the average unemployment duration. However, most research on the determinants of unemployment duration has been confined to an analysis of individual level determinants (Steiner, 1990; Hunt, 1995; Hujer and Schneider, 1996; Steiner, 2001) and the role of the individual employment histories in determining the duration of unemployment (Lüdemann, Wilke and Zhang, 2006; Fitzenberger and Wilke, 2006b). Much less attention has been paid to the regional determinants of the unemployment duration. Most studies only test for additional region-specific effects (Folmer and van Dijke, 1988; Brown and Sessions, 1997; Fahrmeir et al., 2003) and conclude that the regional context is a significant determinant of the individual unemployment duration even after controlling for major individual-specific factors. Other studies only assess the impact of the local unemployment rate or the vacancy-to-unemployment ratio on individual unemployment duration (Lindeboom et al., 1994; Petrongolo, 2001; Haurin and Sridhar, 2003) and typically find the expected prolonging effect of deficient local labour demand on the duration of unemployment. Both of these approaches remain rather incomplete with respect to improving our understanding of the regional factors that prolong or shorten unemployment. We do not know much either about how the institutional context affects individual labour market outcomes. Though there has been a strong interest in the prolonging effect of passive labour market policies such as unemployment benefit entitlements on the duration of individual unemployment (e.g. Carling et al., 1996; Roed and Zhang, 2003; Cockx and Dejemeppe, 2005; Lalive, van Ours and Zweimüller, 2006; Kyyrä and Wilke, 2007), we do not know much about the corresponding impact of other institutional aspects such as

local active labour market policies and local job placement activities. This research gap is particularly surprising in the German context because, among other things, recent labour market reforms emphasise the role of regionally targeted policy mixes and the organisational structure of public employment services. In particular, German policy makers as well as the German public consider a high ratio of job counsellors to unemployed jobseekers as a key to reduce the duration of unemployment.

The objective of this study is therefore to conduct a comprehensive analysis of unemployment duration in Germany. Since the period covered by our data, 2000-2004, falls mainly into the pre-reform institutional setup, we cannot evaluate the success of recent reform efforts. Instead, our regression type analysis aims at exploring the main determinants of the length of unemployment not only among individual characteristics, but also considers the regional and institutional context in which individuals seek employment. By doing so, we provide evidence about the extent to which recent reforms concerning passive labour market measures, regional employment policies and the organisation of public employment services are likely to contribute to a reduction of unemployment duration. For this purpose, our analysis uses a rich set of indicators that capture passive and active labour market policies as well as local economic conditions and job counselling activities. Moreover, we use a new generation of German administrative individual data that allows three main exit states to be identified each of which may be affected quite differently by the regional and institutional context: exits to local regular employment, exits to non-local regular employment via migration and exits to subsidised employment. Previously available data sources did not allow exits to subsidised employment to be distinguished from exits to regular employment. As a consequence, estimated effects of covariates on the duration of unemployment may have been biased if there are heterogeneous effects of covariates on these exit types. In the case of subsidised and regular employment, biases are quite likely because labour market programs typically aim at cushioning unfavourable local labour market conditions. Thus, unfavourable labour market conditions may have an opposing effect on exits to regular and subsidised employment. Similarly, as has been shown in chapter (2), a higher migration hazard may be a response to a deficient local labour demand that lowers the hazard of finding a local job. The paper thus contributes to the literature by disentangling the relevance of individual, regional and institutional factors for exiting unemployment durations to three important exit states.

Moreover, distinguishing between these three exit states also allows for examining whether the availability of active labour market programs (ALMP) affects the probability of finding employment in the regular labour market. Previous research concerning a locking-in effect of ALMP on exits to non-local employment via migration suffered from data limitations as exits to subsidised employment were indistinguishable from exits to regular employment as has been been discussed in chapter (2).

Our findings confirm that for both individual and regional covariates, the impact differs significantly depending on the type of exit. While deficient local labour demand significantly decreases the likelihood of exiting to regular employment in the local area, the likelihood of migration and the likelihood of entering subsidised employment significantly increases. The estimates indicate, however, that individual-level characteristics have a much stronger impact on the duration of unemployment than regional factors. Thus, regional policies may only be a supplementary means of reducing the duration of unemployment. Similarly, local active labour market programs and a higher provision of counselling resources only marginally affect labour market outcomes of unemployed jobseekers and even yield negative labour market outcomes which would be in line with recent results for the Netherlands (van den Berg and van den Klaauw, 2006). Among the regional and institutional factors, our findings indicate that passive labour market policies may have the strongest impact on the duration of unemployment in Germany. This is suggested by extremely low exit hazards to regular employment among individuals with long entitlements to unemployment benefits as well as by major differences in labour market outcomes of unemployed with different income replacement rates.

The structure of our paper is as follows. Section 3.2 gives a detailed description of the unemployment compensation and welfare system and briefly discusses recent labour market reforms. Section 3.3 provides some theoretical underpinning on how job search across multiple labour markets may be affected by regional and institutional factors. Section presents the individual and regional data used in the analysis and discusses the choice of covariates. We then explain the methodological approach before presenting the results in section 3.5. Section 3.6 concludes and discusses the results in light of the recent reforms.

3.2 Institutional context in Germany

Until 2004 the German unemployment compensation system consisted of two main components: unemployment benefits (UB) and unemployment assistance (UA). Unemployment benefits which were paid for a period of up to 32 months, depending on an individual's age and employment history, were equal to 60 % (67%) of the last net income for unemployed individuals without (with) dependent children. The tax-funded and means-tested unemployment assistance was paid indefinitely to individuals who had exhausted their entitlement to unemployment benefit and continued to provide income replacement rates of 53% (57%) for individuals without (with) dependent children. This combination of generous income replacement rates for long-term unemployed and indefinite entitlement length was rather exceptional among the OECD countries. As a consequence, income replacement rates for long-term unemployed in Germany were and still are higher than in many other OECD countries, especially for older unemployed with extended periods of entitlement to UB and for unemployed with low former earnings who receive complementary tax funded social benefits. For this latter group, income replacement rates higher than 70% or even over 100% were thus common practice. From a search-theoretical perspective, high replacement rates raise reservation wages and thus prolong unemployment as the potential net gain from working compared to not working is small (Mortensen, 1990; Rogerson et al., 2005). The institutional design in Germany thus resulted in work disincentives that are considered to be partly responsible for the high share of long-term unemployment in Germany and the considerably higher share of long term unemployment among older people (Fitzenberger and Wilke, 2004) and the low wage unemployed (Fitzenberger and Wilke, 2006b). Moreover, the institutional design has also been associated with a lack of jobs for low-skilled workers in Germany as the social benefit level implies a relatively high minimum wage that is above the productivity level of many low-skilled unemployed. The subsequent empirical analysis of unemployment periods between 2000 and 2004 thus draws specific attention to the unemployment experiences of individuals with low earning capacities.

The "Hartz reforms" introduced between 2002 and 2005 ushered in marked changes in active and passive labour market policies. While the Hartz IV reform that merged social

benefits and unemployment assistance to create the new social benefit[2] (Arbeitslosengeld II) was not implemented before 2005 and is thus not relevant for our analysis, the Hartz I-III reforms already started in 2003. See Jacobi and Kluve (2006) for an extensive overview. These reforms mainly aim at activating the unemployed and increasing the efficiency of employment services and measures. For this purpose, the reform shifts resources from labour market programs aimed at the secondary labour market such as work creation schemes to measures that aim at integrating individuals into the regular labour market (e.g. training, subsidies for regular employment and self-employment). In order to improve the efficiency of allocated resources, programs are targeted more strictly to specific groups of unemployed. Specific reintegration measures are restricted to those who have been assessed to have a fair chance of being reintegrated into the regular labour market, while work creation schemes are targeted to jobseekers with less promising prospects. As a measure to activate job search among unemployed, the reforms also introduced stricter sanction rules in the case of insufficient search efforts, but also offered a new set of programs such as subsidies for people wishing to set up businesses (*Ich-AG*) and subsidies for employers hiring individuals with low productivity levels.

Another key objective of the reforms was the restructuring and modernisation of the Federal Employment Agency (FEA) in order to increase the effectiveness of its placement services. For this purpose, its regional employment agencies introduced a client-oriented New Customer Service Centre (*Kundenzentrum*). An entry zone for customer requests and questions in addition to scheduled appointments for job counselling now prevent long waiting times and increases efficiency. Moreover, computer-based assessments now help in analysing the needs of each customer and thus support tailor-made solutions. These modernisation measures also aimed at reducing the workload of each counsellor in order to improve the quality of job counselling. This new emphasis on job counselling has been facilitated by an increase in the number of job placement counsellors since 2002 of almost 30% and a consequent improvement in the counsellor/customer ratio, i.e. the number of unemployed assisted per placement counsellor.

Another important aspect of the reform concerns the organisation of employment services.

[2]The ALG II provides almost the same level of benefits as former social benefits, while it is below the UA for individuals with high pre-unemployment earnings. The unemployment insurance based UB was basically left untouched.

In contrast to the former hierarchical organisation, far greater responsibility has now been assigned to local employment agencies. Each local employment agency now has to achieve stipulated quantitative goals which are tailored to the specific situation of its regional labour market. For such controlling purposes and the design of regionally tailored policy mixes, the research institute of the Federal Employment Agency, the IAB (*Institut für Arbeitsmarkt- und Berufsforschung*), identified employment agencies with comparable regional conditions. The resulting twelve strategic types of employment agencies range from regional employment agencies with unfavourable labour market conditions in eastern Germany to agencies with favourable and dynamic labour market conditions (Blien et al., 2004). The restructuring of the Federal Employment Agency has therefore resulted in an emphasis on job counselling and efficient placement services as well as an emphasis on regional labour market policies. These internal changes of the FEA were mainly executed by leading international consulting companies who received hundred of millions of euros for their input. An empirical analysis of the institutional features is therefore of high policy interest.

Since the period covered by our data falls mainly into the pre-reform institutional setup of the FEA, we cannot evaluate the success of the restructuring effort. It is, however, possible to obtain empirical evidence about whether one may expect these changes to bring about a strong reduction in unemployment duration. In this respect, our analysis aims at examining the extent to which institutional and regional factors affect labour market outcomes of job-seekers in Germany once individual factors have been taken into account. For this purpose, we explore the impact of passive unemployment measures by looking at the effect of long entitlements to unemployment benefits and by distinguishing between groups of different pre-unemployment wages and thus different income replacement rates. The impact of active labour market programs is captured by explicitly modelling exits to subsidised employment and by examining the impact of the local availability of ALMP on the duration of individual unemployment. In addition, we use a broad number of covariates that capture the regional context and some institutional features such as the counsellor/customer ratio in order to provide some empirical evidence about the extent to which regional employment policies and local job counselling efforts may affect the duration of unemployment in Germany.

3.3 Some theoretical underpinning

Before turning to the empirical approach of our analysis, this section briefly discusses how labour market conditions may affect labour market outcomes after unemployment. For this purpose, consider a framework in which a jobseeker looks for employment in a number of distinct labour markets. In the case of simultaneous job-search across these labour markets[3], the probability of exiting to any of those labour markets can be broken down into the job offer probability and the probability of accepting a job offer in this labour market both of which depend on exogenous market conditions and the endogenous search strategy adopted by the unemployed job searcher. In particular, jobseekers choose reservation wages for each of the distinct markets such that the value of employment at the offered wage is equivalent to the value of continuing the unemployed job search. Moreover, search effort is allocated across the markets so that the marginal value of additional search in each market is equal to the marginal costs of searching the market. While reservation wages affect the job acceptance probability, the allocation of search effort across distinct markets influences the job offer probability. Intuitively, an individual's search strategy should favour finding employment in those labour markets that offer the best working and living[4] conditions. In the case of job search across multiple industries, Fallick (1992) has shown that improving conditions in one labour market - e.g. an increasing job offer probability - raises reservation wages in all markets while at the same time shifting search effort towards the improving market and reducing search effort in all others. As a consequence, changing exogenous conditions affect the hazard of exiting to a specific market not only directly due to, for example, higher job offer probabilities, but also affect these hazards indirectly via the endogenous search strategy of the unemployed job searcher. A similar notion has also been applied to job-search across sectors (Thomas, 1998) and regions (see Damm and Rosholm (2003) as well as chapter (2)).

In our framework, we allow for a local and a non-local labour market and introduce a labour market for subsidised jobs. Exits to non-local employment are likely to constitute only a relatively small but still noticeable share of all exits as migration rates in Germany are low

[3]Alternatively, one may assume some sort of sequential search strategy (Salop, 1973; McCall and McCall, 1987). Accordingly, an unemployed job seeker searches sequentially according to the expected returns from searching a particular market segment.

[4]Since accepting a job in a distant labour market entails a residential mobility, individuals should not only consider job-related characteristics but also the attractiveness of a region as a place of living as has been discussed in chapter (2).

compared with the US, Australia and Canada, but relatively high compared with many other European countries (OECD, 2005). Exits to subsidised employment are likely to constitute a larger part of all exits from unemployment. Compared with other European countries, subsidised employment in Germany is an important part of labour market policy. While spending on active labour market policies in Germany has been around average compared with other European countries, the proportion spent on subsidised employment has been above average in recent years (Martin and Grubb, 2001). Subsidised employment refers to employment in the context of an active labour market program. Such programs mainly encompass subsidised jobs in the secondary labour market, subsidies for regular employment and subsidies for self-employment (see section 3.4 for details). The reforms of recent years have brought about a shift from subsidised jobs in the secondary labour market to the latter two program types (Bundesagentur für Arbeit, 2004). In 2002, more than 200,000 jobseekers entered subsidised jobs in the secondary labour market and more than 350,000 jobseekers received a subsidy for regular employment or self-employment (Bundesagentur für Arbeit, 2002).

Applying the above job search framework across multiple labour markets to our particular setting, jobseekers are simultaneously looking for employment in the market for regular[5] local, regular non-local and subsidised employment. Thus, jobseekers choose the search strategy, i.e. reservation wages and the search effort for each of these markets according to the attractiveness of each of these markets in terms of job availability, offered wages, work and living conditions. Changing conditions in one market may affect all exit-specific hazards via the endogenous search strategy of the jobseeker. In many cases, for example, labour market conditions that favour an exit to local regular employment may have an opposing effect on non-local exits as has been shown in chapter (2). Similarly, subsidised employment is often a means of cushioning unfavourable local labour market conditions. Distinguishing between the three exit states should therefore be quite helpful in understanding how the regional and institutional context affects labour market outcomes of jobseekers in Germany. For this purpose, the empirical analysis considers a number of indicators that capture the exogenous conditions of the local labour market that are discussed in detail in the next section. By

[5]Regular employment can be further differentiated by the number of hours worked or the type of job contract (temporary versus unlimited). However, the data we use does not contain the relevant information such that we pool all types of regular employment.

affecting the search strategy, such conditions not only affect the duration of unemployment, they also affect the probability of making a transition to either local employment, non-local employment or subsidised employment. Other behaviourally distinct and alternative destination states after unemployment that, due to data limitations, are not considered here include exits to self-employment or out of the labour force entirely. Our analysis should therefore be considered as a starting point for improving our understanding of the impact of labour market conditions on the labour market outcomes of unemployment.

3.4 Data

This section describes how we select the sample and covariates for our analysis. We use individual data merged from several administrative registers which is then combined with regional data from various sources.

Individual data The Sample of the Integrated Employment Biographies V.1 (IEBS) of the Research Data Centre (*Forschungsdatenzentrum*) of the Federal Employment Agency (FEA) is a new data set which was released in 2005. See Hummel et al. (2005) for a detailed description of the data. It is a 2.2% sample containing about 1.4 million individuals in the period 1992-2004. It comprises high quality information about employment periods that have been subject to social insurance contributions and thus excludes civil servants and self-employed individuals. The sample also contains information on the receipt of unemployment compensation from the FEA. In addition, the data set provides information about participation in active labour market programs for the period 2000-2004. One of the major drawbacks of the data is that it only partially identifies the true unemployment period. This is because there are unobserved periods in the employment trajectories whenever an individual is neither a socially insured employee nor receives unemployment compensation, nor participates in any active labour market program. As a consequence, some parts of the individual employment trajectory may not be observed so that it is necessary to use proxies for the true unemployment period, see e.g. Fitzenberger and Wilke (2004) and Lee and Wilke (2005) for this problem. In the analysis of this paper we use the following proxy for the true unemployment duration:

- **Unemployment with permanent income transfers (UPIT)** is a lower bound of the true unemployment period that defines unemployment as a continued period of transfer receipt. Gaps between transfer receipt and the beginning of a new employment period need to be less than four weeks. Thus, UPIT excludes periods of unemployment without receipt of UB or UA from the FEA.

Unfortunately, there is no exact way of telling whether this unemployment proxy more closely resembles the true length of unemployment than competing proxies. Moreover, the data does not contain all necessary information to identify unemployed social benefit recipients, a group that we are specifically interested in as discussed in section 3.2. Entitlements to complementary social benefits depend on pre-unemployment earnings but also on the number of dependent children as well as financial resources (e.g. spouse income, private savings). Since we do not observe enough information in the IEBS about the household context nor about its financial resources, no exact identification of unemployed social benefit recipients is possible. Individuals with relatively low pre-unemployment wages, however, are more likely to receive additional social benefits. We therefore compare unemployment periods of social benefit recipients which are contained in the Social Benefit Statistics (Sozialhilfestatistik, SH-Stat)[6] with unemployment spells in the IEBS for individuals of different pre-unemployment earnings in order to choose an income threshold below which unemployed in the IEBS are similar to unemployed social benefit recipients. Moreover, we conduct this comparison not only for the above UPIT proxy of unemployment in the IEBS, but also for a wider proxy which also adds nonemployment periods to the unemployment duration. A comparison of the corresponding distributions of unemployment duration suggests that the UPIT proxy for individuals with pre-unemployment gross earnings of less than 60 euros per day better represents the group of unemployed social benefit recipients than the competing unemployment proxy or other income thresholds. A daily gross wage of 60 euros closely corresponds to the lowest wage quintile for full-time employees in western Germany and to the lowest two wage quintiles for full-time employees in eastern Germany. Applying the same income threshold for both parts of Germany may appear somewhat crude. Robustness checks using, for exam-

[6]The use of the SHStat was confined to the research project *Evaluation of the experimentation clause §6c SGB II* which was funded by the German Ministry of Labour and Social Affairs. No scientific use file exists for this unique data set such that apart from the comparison of both data sets, no further analysis could be conducted. For more details on the comparison of the data sets see Arntz et al. (2006b).

ple, unemployed in the lowest wage quintile for both parts of Germany, did not significantly change the results. We therefore decided to apply the UPIT definition in the subsequent empirical analysis and stick to the chosen threshold of 60 euros daily gross earnings to distinguish individuals of low-earning capacities from individuals with higher earning capacities. Individuals above this threshold are less likely to receive additional social benefits and should thus have different unemployment experiences than their low-earning counterparts.

For all UPIT unemployment spells, we observe the exit state if the spell is not right-censored due to the end of the observation period and if the unemployed continuously receives income transfers from the FEA. As discussed in the theoretical section, we distinguish between local regular employment, non-local regular employment (migration) and subsidised employment. We define migration as movements between non-adjacent labour market regions (*Arbeitsmarktregionen*). The 227 labour market regions (LMRs) in Germany comprise typical daily commuting ranges such that for the majority of individuals the workplace is located within the LMR. Finding employment in a non-adjacent LMR therefore usually necessitates residential mobility. We refer to subsidised employment whenever an individual exits to socially insured employment or self-employment in the context of an active labour market program. Such programs mainly encompass subsidised jobs in the secondary labour market (*ABM, SAM*), subsidies for regular employment (*Eingliederungszuschüsse, Beschäftigungshilfen*) and subsidies for self-employment (*Ich-AG, Überbrückungsgeld*), but also contain more extensive training programs (*FbW*) if these programs count as socially insured employment. Table 3.1 describes the composition of all exits to subsidised employment observed in the IEBS for UPIT spells starting between 2000 and 2002. For the analysis, we decided to pool all forms of subsidised employment because robustness checks for distinguishing between certain types of programs did not yield noteworthy differences compared to pooling all programs.

We restrict our analysis to unemployment periods starting in the period 2000-2002. This is because information on periods of subsidised employment is not available before 2000. Since we are able to observe information about unemployment up to 2004 while exits to employment are only observable up to the end of 2003, we decided to exclude spells starting in 2003. This reduces the amount of right censoring in the data and ensures a minimum observation period of one year. Table 3.2 shows the sample sizes and exit states when

applying the UPIT definition and distinguishing individuals by their earning capacities. We also distinguish by gender and marital status as these characteristics are important determinants of individual labour market outcomes.

Table 3.1: Composition of exits to subsidised employment, IEBS, 2000-2002

Subsidy for ...	**Number**	%	**% of all exits**
... employment in secondary market	10,391	31.0	7.7
... regular employment	9,643	28.7	7.1
... self-employment	9,001	26.8	6.7
... training measure	2,146	6.4	1.6
... other programs[a]	2,379	7.1	1.8
Total subsidised employment	33,560	100.0	24.9

[a] This category refers to a mix of programs that can be autonomously designed by each employment agency. As an example, these measures include subsidies for entering vocational training or a premium for extending working hours of an existing job (Bundesagentur für Arbeit, 2002).

Table 3.2 shows that individuals with low pre-unemployment wages are more likely to exit to subsidised employment and less likely to migrate than jobseekers with higher pre-unemployment earnings. Moreover, the median unemployment duration is significantly longer for individuals with low pre-unemployment wages, a finding that is in line with the expectations that the institutional framework creates disincentives for individuals with low earning capacities to take up a job. Table 3.2 also indicates differences by gender and marital status. Singles are geographically more mobile than their married counterparts, a finding that is consistent with the migration literature regarding higher migration costs for married people with children (see Ghatak et al., 1996). Differences between female and male singles, however, are very small. Since estimation results for single males and females proved to be very similar, we decided to pool male and female singles in the subsequent analysis. By contrast, results for married individuals strongly differ by gender. Married women have by far the longest median unemployment duration and the lowest exit rates. This probably reflects the looser labour force attachment of married women. Moreover, the extremely low migration rates among married women may reflect the fact that women are more likely to be tied to the local area if the male breadwinner is employed locally. Due to these particularities of labour market decision of married women, we decided to restrict the analysis to married

males and singles only and differentiate these groups by their pre-unemployment earnings.

Table 3.2: Unemployment duration and exit types by gender, marital status and earning capacity, IEBS, 2000-2002

	Low Wage[a]		**Higher wage**[a]	
	Men	**Women**	**Men**	**Women**
Singles				
% exit to				
local employment	48.1 (70.3)	51.3 (74.0)	57.9 (74.5)	53.5 (70.5)
non-local employment	6.4 (9.4)	5.6 (8.1)	9.1 (11.7)	10.3 (13.6)
subsidised employment	13.9 (20.3)	12.4 (17.9)	10.7 (13.8)	12.1 (15.9)
all exits	68.4 (100.0)	69.3 (100.0)	77.7 (100.0)	75.9 (100.0)
Unemployment spells				
Median duration (days)	138	146	107	123
Number of spells	43,528	31,206	20,849	7,319
Married				
% exit to				
local employment	47.5 (64.9)	44.6 (60.9)	58.7 (73.3)	45.9 (72.6)
non-local employment	5.8 (7.9)	2.8 (3.4)	8.0 (10.0)	4.6 (7.3)
subsidised employment	19.9 (27.2)	14.7 (20.1)	13.4 (16.7)	12.7 (20.1)
all exits	73.2 (100.0)	62.1 (100.0)	80.1 (100.0)	63.2 (100.0)
Unemployment spells				
Median duration (days)	176	238	116	194
Number of spells	28,018	31,088	23,620	5,483

[a] Low wages refers to individuals with pre-unemployment daily gross wages of less than 60 euros, while higher wages denotes individuals with pre-unemployment earnings above this threshold.

Individual-level covariates for the econometric analysis that are contained in the IEBS are age, education and a number of indicators of an individual's employment history such as previous unemployment experiences, previous participation in active labour market programs and previous commuting status. These covariates are chosen to capture differences in job-finding chances and migration costs that are relevant for the labour market outcomes of jobseekers. In addition, the analysis also includes the maximum length of unemployment benefit receipt at the beginning of unemployment. Long entitlement periods may prolong unemployment and should thus be included in the analysis. As previously discussed, entitlements to unemployment benefits increase with age and job tenure within some claim period. Since actual entitlements are not observed for many individuals in the IEBS, the missing

information has been imputed based on known information concerning age and tenure in the previous job.[7] Summary statistics of the samples can be found in Appendix A.

Regional aggregate data We use a broad number of regional indicators which are mainly provided by the two largest German data producers: the Federal Employment Agency and the Federal Statistical Bureau (FSB). The FEA data is coded at the level of employment agency districts and contains monthly information about labour market tightness (e.g. vacancies, jobseekers, degree of long term unemployment), the extent and structure of local labour market programs and the organisation of the local employment agency (e.g. number of staff). The FSB data contains yearly county level information about the population structure (e.g. age, education), the type of region (urban vs. rural), its infrastructure and industrial structure. There are 180 employment agency regions and 440 counties in Germany. Since the IEBS identifies both the county and the employment agency region of the workplace, we decided to merge the data at the regional level for which they are originally provided. This avoids problems associated with interpolating attributes from one of these regional classifications to the other as has been discussed in chapter (1).

The FSB and the FEA data together contain more than 100 regional indicators, a full list of which is included in Arntz et al. (2006b). These are far too many regional covariates for an econometric analysis and there is a high degree of correlation among several of these regional indicators. Thus, as a first step we used a combination of cluster and factor analysis to identify indicators that contain very similar information. In a next step, we decided to compress the regional information further by grouping the remaining regional indicators according to economically reasonable groups that cover major regional factors that are likely to affect unemployment durations and the labour market state after unemployment as discussed in the theoretical framework. In particular, we create five groups and select up to five indicators as their representatives such that the correlation among the representatives is minimised. As a consequence, the chosen representatives proxy for their group of interest in the econometric analysis so that estimated coefficients reflect effects of the group they represent. Table 3.3 shows a description and summary statistics of all regional indicators.

[7] Gaps between two employment spells of less than one month have been considered as one continuous job. The resulting measure is likely to be a lower bound of the true entitlements because remaining entitlements from a previous claim of unemployment benefits are not taken into account.

There seems to be enough regional variation in most of the indicators that describe the regional labour market situation to identify the effect of regional covariates on labour market outcomes. For the subsequent econometric analysis, we standardised all continuous regional variables to ease comparability of estimation results.

The first group of indicators characterises local labour **demand and supply conditions**, i.e. local job availability. The local unemployment rate may be considered as an indicator of deficient local labour demand. In addition, the change in the unemployment rate compared to the previous year conveys information about the development of the local imbalance of labour supply and demand. In regions with an excess supply of labour, the probability of receiving a job-offer should be reduced. As a reaction, reservation wages in all labour markets decrease since jobseekers become less choosy and search effort shifts from the local to alternative markets. This implies a decrease in the number of local jobs found and an increasing hazard of finding a non-local or a subsidised job. Increasing exits to subsidised employment in the case of an excess supply of labour may also reflect an increasing availability of subsidised employment because labour market programs are often used to cushion unfavourable labour market conditions.

Another important determinant of unemployment duration might be local **economic performance** since well-performing and dynamic regions should offer a higher expected lifetime income and should thus attract search effort to the local market while non-local and subsidised employment should become less attractive. Well-performing and economically growing regions should be characterised by a high and growing GDP per head as well as by a high level of newly established businesses which we all include as covariates.

Apart from economic conditions of the locality, its **social structure** may also shape individual labour market behaviour. In particular, individuals may have "lower incentives to work where peers are also unemployed ... and a view of joblessness as unproblematic within a context of lowered aspirations, ..." (Ritchie et al., 2005:3). In Germany, discouraging social contexts might be found in old industrial regions which have experienced massive deindustrialisation in recent decades and a subsequent rise in long-term unemployment. We thus decided to include indicators such as the level of long-term unemployment and the

average schooling level in the region to control for different social contexts.[8]

In addition, we use information about the **institutional organisation** of the local employment agency. As discussed in section 3.2, there has been an increase in the number of job placement counsellors of around 30% during the period of observations. This politically motivated increase in the counsellor/customer ratio, i.e. the ratio between placement officers per jobseeker, provides some exogenous variation to identify the effect of an increasing level of job counselling. We hypothesise that a higher counsellor/customer ratio shortens the duration of unemployment, especially by accelerating exits to regular employment. We also include indicators of the local availability of labour market programs. As discussed in section 3.2, there has been a shift from measures aiming at the secondary labour market to programs that aim at integrating individuals into the regular labour market. Moreover, individuals with different employment prospects may be affected differently depending on the type of program due to the increased profiling efforts of employment agencies. We therefore distinguish between programs with a focus on the regular labour market such as training measures (*FbW*), programs targeted to young unemployed (*JUMP*) and subsidies for regular employment or self-employment (*Übergangsgeld, Eingliederungszuschuss, Beschäftigungshilfe*) and programs with a focus on the secondary market such as work creation schemes (*ABM, SAM*) and include the share of unemployed participating in these program types, respectively.[9] While exits to subsidised employment should be positively affected by the level of offered programs, the hazard of leaving the region may be negatively affected if such programs offer a substitute for leaving the region. In this case, jobseekers may reduce search effort in non-local labour markets. An extensive local availability of ALMP may thus result in a regional locking-in effect as has been discussed in the Scandinavian literature (e.g. Westerlund, 1998; Fredriksson, 1999).

[8]Both of these indicators are highly correlated to the share of social benefit recipients in the region which we therefore decided to leave out. The two remaining indicators should capture this regional characteristic.

[9]Further differentiating the program types is problematic as we often found a high degree of correlation between similar program types.

Table 3.3: Description and summary statistics of regional covariates, 2000-2003

No.	Group	Indicator	Source[a]	Mean	SD
I	**Demand / supply**	• Unemployment rate	A	10.7	5.0
		• Change of U rate on last year (pp)	A	0.04	0.71
II	**Economic performance**	• GDP per employee (euros)	B	49818.5	8891.4
		• Change of GDP per employee 1995-2000 (%)	B	2.4	1.6
		• New businesses set up per 1,000 residents	C	13.8	2.7
III	**Social structure**	• Share of long-term U[b] (%)	A	33.7	6.7
		• Avg. years of schooling	B	14.6	0.13
IV	**Institutional organization**	• Placement counsellor/100 unemployed	A	0.24	0.05
		• Share of unemployed in ALMP[b] with focus on:			
		- regular labour market - $ALMP_R$ (%)	A	17.6	3.9
		- secondary labour market - $ALMP_S$ (%)	A	5.9	5.8
V	**Structural indicators**	• Driving time to next large city (min.)	B	104.8	53.8
		• Child care places per child	B	0.62	0.12
		• University present (Yes=1)	B	0.36	0.48
		• Type of the region (reference: sub-urban):			
		- rural	B	0.29	0.45
		- urban	B	0.39	0.49
		• Unemployment turnover per 1,000 employees	A	32.2	11.1

[a] A: *Bundesagentur für Arbeit* (FEA); B: *Statistisches Bundesamt* (FSB); C: Institut für Mittelstandsforschung, Bonn.
[b] ALMP=Active labour market programs; U=Unemployment.

Finally, we include several **structural indicators** to characterise the type of region. In particular, we include a population density related classification to distinguish between rural and urban regions. Moreover, we use driving distance to the next higher level city as a proxy for the degree of remoteness of a region. Both of these characteristics affect the availability and the accessibility of employment and may thus change an individual's search behaviour. Moreover, an individual's search behaviour may also be affected because these characteristics to some extent reflect the attractiveness of a region as a place of living (e.g. availability of consumer amenities, disamenities such as pollution). We also control for three specific other regional characteristics. Regions with a high level of seasonal work, proxied for by the flow in and out of unemployment, may be characterised by a large share of short unemployment spells. Secondly, the local existence of third level institutions may affect the composition of the available workforce and thus the competition for certain jobs. Finally, we include the local availability of child care support in order to test whether the public infrastructure affects unemployment experiences of jobseekers with children. The availability of kindergarten or nursery school might reduce the opportunity costs of local employment and thus accelerate exits to local employment.

3.5 Methodological issues

Let $F(t)$ be the unemployment duration distribution, where t is the duration of unemployment. The hazard rate, $h(t) = \{\partial F(t)/\partial t\}/(1 - F(t))$, is an intuitive way of formalising transitions from unemployment to employment. In our econometric analysis we use a hazard rate model to investigate the effect of various covariates $x = \{x_1, x_2\}$ on the distribution of unemployment, where x_1 denotes the set of individual characteristics such as demographic and socioeconomic characteristics, work history variables and firm-level variables, while x_2 contains all remaining regional indicators. In particular, we assume the different exit states to be independent conditional on all covariates[10] and estimate a competing-risk Cox proportional hazard model

$$h_j(t|x) = \lambda_j(t)exp(\alpha_j x_1 + \beta_j x_2),$$

[10]If this assumption is violated, estimates for the independent competing-risk model may be biased. Since this assumption only needs to hold conditional on all covariates and the analysis includes a rich set of covariates, the assumption may be justified.

where h_j denotes the exit-specific hazard rate for the three destination states local regular employment, subsidised employment and non-local employment, i.e. migration. λ_j is the destination-specific baseline hazard.

There are three major sources of biases that have to be addressed when using this approach. First of all, there may be biases from unobserved individual heterogeneity. As suggested by Meyer (1990), however, unobserved individual heterogeneity may not have much of an effect if there is a flexible baseline hazard that partly absorbs this heterogeneity. Secondly, there may be a simultaneity issue of the regional covariates if an explanatory variable x_t depends on observed exits from unemployment in period $\tau \geq t$. Due to the discrete nature of the data, this may be the case for certain contemporaneous labour market characteristics such as the unemployment rate, the accommodation of local active labour market programs or the counsellor/customer ratio. For this reason, all labour market related information provided by the FEA for which such an issue may arise have been calculated as the average value for the 12 months preceding the start of unemployment. Estimation results may, however, still be biased due to unobserved regional heterogeneity. In the literature, this problem has been addressed by stratification (see Ridder and Tunali, 1999). When stratifying according to regional labour markets, separate baseline hazards are estimated for each regional labour market. This approach resembles the well-known fixed effects approach and thus controls for unobserved heterogeneity at the level of regional labour markets. Unfortunately, our data is limited to a relatively short time span. Thus, a stratified estimation approach turned out to be infeasible since, in this case, identification rests on time variation. We are nonetheless fairly confident that biases from omitted regional characteristics may be negligible due to the rich account of regional covariates used in the analysis.

As has been discussed by Thomas (1996), in a competing-risk duration analysis, the estimated parameter vector for a particular destination state may not be interpreted as the effect on the duration until exiting to this state. Instead, this effect as well as the effect on the probability of leaving to a particular state depends on parameter vectors for all states. In particular, define the conditional cumulative probability of exiting to state j until t as

$$\Pi_j(t|x) = \int_0^t h_j(t|x)(1 - G(t|x))dt$$

with $h_j(s)$ as the exit hazard to state j and $(1 - G(s))$ as the overall survival probability that takes account of all exit options. In our empirical analysis we evaluate the estimates at

$x_l \in \{\bar{x}_l, 0\}$, where we choose the average values of all individual level variables ($x_1 = \bar{x}_1$) and we choose zero for the regional variables ($x_2 = 0$).[11] We estimate the probability of exiting to state j as the duration elapses one year, i.e. $\Pi_j(365|x)$. We compute the marginal effects $\partial\Pi_j(365|x)/\partial x_k$ as the marginal change of the cumulative probability of exiting to state j during the first year if one regressor x_k changes. This outcome is of particular political interest because in official statistics long-term unemployment starts after one year of unemployment. Thus, our marginal effects correspond to the change in probability of becoming long-term unemployed that is due to a marginal increase of covariate k.[12] Based on 500 samples, we estimate the standard error of the conditional marginal effect bootstrap distribution. Assuming that standard errors are distributed normally, we then determine the significance level of the estimated marginal effects.

3.6 Results

Tables 3.4 and 3.5 present the estimated conditional marginal effects for singles and married men with low and higher pre-unemployment wages. We generally find that the individual work-history seems to be the driving force behind the duration of unemployment, a result that is similar to Lüdemann et al. (2006) and Fitzenberger and Wilke (2006b) who use data without information on subsidised employment and on migration. Compared to the impact of individual characteristics, regional disparities only marginally affect the length of unemployment periods in Germany as has also been suggested by the analysis in the preceding chapter that uses data without information on subsidised employment. Thus, although some regional factors significantly affect both the unemployment duration and the likelihood of ending up in a specific destination state, our results suggest that the recent emphasis on regionally tailored policy mixes and job placement activities is unlikely to bring about a substantial reduction in the length of unemployment in Germany. In what follows

[11] Due to the standardisation of all continuous regional variables, this corresponds to the sample mean value for these covariates and to the reference category of the regional dummy variables.

[12] Since $\Pi_j(t|x)$ has the properties of a distribution function, one may define the conditional marginal quantile effect at quantile q as $\partial\hat{t}_j(q|x)/\partial x_k = \partial\Pi_j^{-1}(q|x)/\partial x_k$ as an alternative marginal effect. Since the underlying unemployment duration distribution is defective, $\Pi_j^{-1}(q|x)$ does not exist for the upper quantiles so that $0 \leq \Pi_j(t|x) \leq \bar{q}_{|x} \leq 1$. Moreover, the maximum quantile for which this marginal effect can be identified varies by covariate and destination state, i.e. $\bar{q}_{jk}$. For this reason, we decided to report the marginal effect on the cumulative probability $\Pi_j(365|x)$ only. Marginal quantile effects are available from the authors upon request.

we present a detailed discussion of the estimation results for the individual-specific covariates before turning to the regional covariates. In line with the finding that regional covariates have only a limited impact on individual labour market outcomes, we also find only few general and robust result patterns across the four sub-groups. A detailed discussion of each single effect thus seems an infeasible approach. Instead, we only focus on the most important results for each group of regional covariates and point to the most interesting and robust differences across the sub-samples.

Demographic and socioeconomic characteristics Several socioeconomic variables significantly affect the duration of unemployment[13], but only few of them have a strong effect. Among the most important for all exit states of singles and married men alike is age. Generally, the probability of finding regular employment in either the local or the non-local area decreases with age. For low-earning individuals aged 46-56, the reduced probability of finding regular employment is partially compensated by a higher probability of entering subsidised employment. Those aged 56 and above experience lower exit probabilities to all exit types and thus stay unemployed significantly longer than their younger counterparts. Once having controlled for different levels of pre-unemployment wages, the effect of educational attainment on the duration of unemployment is rather limited. Instead, the educational degree rather has some impact on the observed exit state after unemployment. In particular, a university degree markedly reduces the probability of finding local regular employment, but increases the likelihood of entering subsidised employment or migration if they are single. The resulting probability of leaving unemployment to any of these exit states within one year thus tends to be slightly lower only than for their unskilled counterparts. Lower pre-unemployment earnings associated with higher income replacement rates rather than the observed educational degree thus seem to be able to explain the high share of long term unemployment among the unskilled in Germany. This finding also confirms our approach of stratifying the sample with respect to the pre-unemployment wage level.

Work history Characteristics associated with the work history have the strongest influence on the unemployment duration distribution and effects are typically similar for all sam-

[13]When the effect is similar for all destinations we simply use the notion *unemployment* or *unemployment duration*.

ples. In particular, long entitlement periods for unemployment benefits (UB) are associated with a much lower likelihood of finding regular employment either in the local or the non-local region, especially among individuals with higher pre-unemployment wages. One reason for why UB entitlements have a weaker effect on individuals with low pre-unemployment wages may be that these individuals often receive complementary social benefits. In this case, prolonging unemployment benefits does not result in additional transfer payments as the income replacement rate is the same irrespective of whether receiving UB or unemployment assistance (UA). For those with higher pre-unemployment earnings, longer UB entitlements may instead be quite beneficial as income replacement rates for this group should be higher for UB than for UA. In line with our findings, the disincentive effect from longer UB entitlements should be stronger for individuals with higher pre-unemployment earnings. Another interesting finding concerns extremely long UB entitlements of more than 2 years which apply only to older unemployed who already have much lower exit probabilities than their younger counterparts. These individuals basically never leave unemployment again but use their long entitlements to unemployment benefits as a means of early retirement.

We also obtain strong results if an unemployed person was already subsidised by the local employment agency before the start of the current unemployment period. If these individuals slip back into unemployment they have a very low transition probability to either local or non-local regular employment. Instead, a high percentage of these individuals ends up in another subsidised employment period. With the new generation of individual administrative data we are therefore able to identify what is typically called a "career of labour market measures".[14] Our results therefore suggest that both passive and active labour market measures are strongly associated with negative individual labour market outcomes. We do not, however, read this as a pure causal relationship because these results may partially be driven by unobserved factors such as a selection of immobile unemployed into subsidised employment or long entitlements to UB. Since entitlements to UB depend on a positive employment history, a negative selection may be less likely, however, to explain the negative labour market outcomes of individuals with long entitlements to unemployment benefits.

[14]We also made estimations in which we distinguished between several types of employment subsidies offered by the employment agencies. Surprisingly, the results patterns are similar even for subsidised artificial jobs and temporary subsidies of regular employment which have a very purpose. For this we decided to report the pooled results only.

Furthermore, we identify several factors that increase exit probabilities among the unemployed. Individuals who have previously been recalled by their former employer, have much shorter unemployment periods due to faster local exits. Moreover, this group is less likely to be subsidised or to migrate, and this suggests that recalls are related to seasonal unemployment and temporary lay-offs. Being in minor employment[15] at the beginning of the unemployment period considerably increases local job finding and reduces the likelihood of migration in many cases. Working in a minor employment signals some labour force attachment, but may also increase the attachment to the local area. Individuals who commuted to their last job have lower local but higher non-local employment probabilities. This may capture both a higher propensity to migrate as well as a higher propensity to commute very long distances.

Supply and demand conditions As expected, a deficient local labour demand as reflected in high and increasing unemployment levels tends to reduce the probability of finding employment locally within one year among all groups. Among single persons with higher pre-unemployment wages, this detrimental effect on the duration of unemployment is partially offset by higher migration levels while their married counterparts increasingly enter subsidised employment in regions with an excess supply of labour. For individuals with low pre-unemployment wages, such counteracting effects are even smaller or absent so that this group appears to be more reliant on local labour market conditions. Similar to the findings in chapter (2), the results thus indicate some heterogeneity among different groups of unemployed individuals with respect to the responsiveness to regional labour market conditions. Moreover, the fact that significantly higher migration levels can only be found for well-earning singles in response to deteriorating unemployment rates during the previous five years suggests that such counteracting forces tend to be slow, a result that is in line with slow adjustment processes after region-specific shocks (Decressin and Fatàs, 1995; Möller, 1995). Thus, deteriorating local labour demand conditions seem to result in a growing share of long-term unemployed at least in the medium-term.

[15] An employment on a salary of less than 400 euros per month and with exemption from social security contributions.

Table 3.4: Marginal effects in pp on the conditional cumulative probability of exiting to local, subsidised or non-local employment, Singles, IEBS, 2000-2002

	Low wage			High wage		
Variable	local	subsidised	non-local	local	subsidised	non-local
Individual characteristics						
Female	7.6**	-0.8†	0.6†	2.9*	-0.5	2.1**
Age < 26	16.7**	-0.9**	1.2**	11.7**	-2.2**	0.0
Age 26-35	3.7**	0.4†	1.1**	4.2**	0.7†	1.0**
Age 46-56	-9.6**	1.5**	-0.9**	-4.5**	0.2	-0.2
Age > 56	-22.2**	-0.8	-1.7**	-12.2**	-5.4**	-0.1
Unskilled	-3.0**	0.5**	0.2	-2.8**	0.0	1.4**
University degree	-2.6	3.0**	3.6**	-9.8**	3.2**	1.6**
Foreign born	-1.4	-1.0*	0.6	-3.8**	-1.3	-1.3**
Female foreign born	-2.1	-1.9*	-1.4**	-1.1	1.5	-1.4
Children	-2.1**	0.7**	-0.1	-1.8†	0.4	0.4
Children & female	-4.7**	-0.2	-2.2**	-1.7	1.1	-2.6**
Minor job	9.2**	0.0	-0.4†	3.9	1.5	-3.4**
Spell starts in winter	4.7**	0.8**	0.2	7.1**	-1.0**	0.2
Previous employment history						
Part time	-5.3**	-0.9**	0.0	5.0**	-1.2†	2.4**
Lower/upper wage‡	-2.7**	1.1**	-0.5**	-5.2**	1.6**	4.5**
Lower/upper wage† & female	0.8	-0.2	0.6*	2.2	0.7	-0.8*
Construction	8.5**	-1.2**	-0.8**	10.3**	-1.9†	-2.2**
Trade and Food Ind.	7.7**	-0.6*	-0.1	3.8**	0.8	-2.0**
Services/Public sector	3.8**	-0.5**	0.6**	-3.4**	1.7**	-0.5*
Previously recalled	15.7**	-1.7**	-0.4*	14.8**	-4.6**	0.4
Previously unemployed	1.3**	1.0**	-0.4**	1.9**	-0.4	0.0
Large firm	-2.7**	-1.2**	0.2	-5.9**	-1.6**	0.7*
BE > 6 − 12 mths	3.7**	0.3	0.1	-2.7**	2.2**	-1.2**
BE > 12 − 18 mths	-1.0	0.5	0.2	-9.7**	2.5*	-1.5**
BE > 18 − 24 mths	-3.7	0.9	-1.1†	-15.1**	4.0**	-2.9**
BE > 24 mths	-21.0**	-1.7*	-1.9**	-28.2**	1.6	-4.7**

Continued on next page...

... table 3.4 continued

ALMP measure	-31.1**	7.2**	-2.0**	-25.2**	12.7**	-1.8**
Commuter	-4.6**	0.3†	5.6 **	-11.8**	0.7†	8.0**
Regional characteristics						
Unemployment rate (UR)	-2.3*	0.6	0.0	-0.5	-1.8*	1.0
Change in UR 1995-2000	-1.5**	0.5**	0.2	-3.3**	0.7*	0.7**
GDP per head	-0.5	-0.2	0.2	-0.7	0.3	0.4*
Change of GDP 1995-2000	0.7**	0.1	-0.2**	0.3	-0.5**	-0.2*
Rate of business set ups	1.1**	-0.2†	0.0	-0.2	0.2	0.3†
Share of long-term U	-1.9**	-0.4*	0.2	-3.2**	0.7*	0.9**
Avg. yrs of schooling	1.8**	-0.2	-0.5**	0.6	-0.3	-0.4*
Placement counsellor per U	0.2	-0.4*	0.4**	-1.2*	-0.2	1.6**
Share of U in $ALMP_R$	0.1	0.5**	-0.1	-0.1	0.6†	-0.1
Share of U in $ALMP_S$	-1.4**	0.9**	0.1	-1.5*	1.1**	0.0
Driving time to higher level city	0.8**	-0.2	-0.2†	-0.2	0.1	0.2
Child care places & child	0.3	-0.3†	-0.4**	0.4	0.4	-0.1
University present	-2.6**	0.1	0.4 **	-1.8**	-0.3	1.0**
Saisonal unemployment	-0.2	0.2	0.1	-2.5†	1.8†	0.0
Rural region	-1.8**	0.0	0.2	1.6†	-1.4**	0.2
Urban region	-0.1	1.0**	-0.9**	1.6*	-0.6	-0.4
West & 2000	8.6**	0.5	1.4**	9.9**	-1.1†	1.0*
West & 2000 & female	0.4	0.0	0.0	1.3	-0.3	-1.0*
West & 2001	2.5**	1.2*	0.3	0.1	0.5	0.8
West & 2001 & female	2.0†	0.2	0.5	1.7	-0.8	0.0
East & 2000	-3.8**	2.0*	3.7**	-7.8**	2.3	11.3**
East & 2000 & female	-6.5**	0.0	-0.7*	1.6	-2.1†	-2.7**
East & 2001	-2.0	0.5	3.6**	-5.4**	1.7	6.8**
East & 2001 & female	-6.0**	0.3	-1.0**	-1.8	-1.3	-1.5*
East & 2002	0.1	0.0	2.6**	-2.3	1.5	3.1**
East & 2002 & female	-6.6**	0.0	-1.1**	-0.2	0.5	-1.5**

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Note: Low wage sample refers to individuals with pre-unemployment daily gross wages ≤ 60 euros. BE=Benefit entitlements; ALMP=Active labour market program with focus on regular (R) or secondary (S) employment; U=Unemployment.

‡: Daily pre-unemployment wages are in the lowest (highest) wage quartile for the low (higher) wage sample.

Economic performance Indicators that proxy for the local economic performance do not show any robust pattern across the different groups of unemployed. The only exception is the strong and positive effect of the setting up of new local businesses on the likelihood of finding local employment within one year for low-earning individuals. One reason for this positive effect may be that new firms tend to offer precarious jobs which are a more relevant type of employment for individuals at the fringe of the labour market. Apart from this noteworthy effect, the effects of other local indicators of economic performance are negligible. We therefore conclude that local economic performance does not seem to be an important determinant of labour market outcomes for jobseekers in Germany, one explanation of which may be that due to central wage bargaining regional productivity levels as reflected in local GDP do not translate into behaviourally relevant interregional wage differences. In fact, gaps between regional productivity and wage levels may be one reason for persistent interregional employment disparities in Germany.

Social structure While the effect of the average schooling level shows no clear result pattern across the different sub-samples, the share of long-term unemployment (LTU) seems to mainly confirm the negative effects of a discouraging social context. As expected, a high share of LTU significantly prolongs the duration of unemployment as the strong decrease in local exits is only partially offset by increasing exits to non-local and subsidised employment. We are however careful in reading this as a pure discouraging effect as this effect to some extent may also reflect that regions with a high share of LTU are characterised by bad job-finding conditions that prolong unemployment.[16] On the other hand, controlling for the unemployment rate during the last year and the change in unemployment during the last five years suggests that the remaining variation in LTU mainly stems from an unfavourable composition of individuals who tend to experience prolonged unemployment periods. In this case, the effect may indeed rather reflect a discouraging effect on the search activities of jobseekers, although there is no evidence that overall search effort is reduced as, at least for some individuals, the probability of migration and subsidised employment increases.

[16]We do not think that there is a severe problem with reverse causality because we use the lagged share of LTU and control for a rich set of regional characteristics that may be related to the share of LTU.

Table 3.5: Marginal effects in pp on the conditional cumulative probability of exiting to local, subsidised or non-local employment, Married Men, IEBS, 2000-2002

	Low wage			High wage		
Variable	local	subsidised	non-local	local	subsidised	non-local
Individual characteristics						
Age < 26	10.2**	0.0	0.8†	1.9	0.1	-0.2
Age 26-35	2.1 **	0.5	0.5*	2.3**	-0.3	-0.1
Age 46-56	-8.3**	2.8**	-0.8**	-5.7**	0.0	0.0
Age > 56	-22.1**	0.5	-2.2**	-20.6**	-5.6**	-1.2**
Unskilled	-3.4**	1.1**	0.0	-1.5*	-1.5**	0.1
University degree	-7.0**	3.6**	0.1	-10.0**	2.5**	0.4
Foreign born	-2.0**	-1.5**	0.3	-5.7**	-2.8**	0.8*
Children	-0.5	0.2	-0.1	-0.6	0.0	-0.2
Minor job	13.9**	-0.4	-0.6	9.0**	1.0	-2.9**
Spell starts in winter	10.0**	0.4	0.2	11.3**	-2.7**	0.1
Previous employment history						
Part time	-4.6**	-0.9*	-0.4		n/a	
Lower/upper wage‡	-8.2**	0.8*	-0.5**	-5.7**	2.8**	2.7**
Construction	16.6**	-1.8**	-0.3	16.0**	-3.4**	-2.8**
Trade and Food Ind.	12.0**	-1.0*	0.2	6.5**	-0.2	-1.0**
Services/Public sector	6.9**	-0.6†	0.8**	-0.4	1.9**	-0.1
Recall	12.3**	-2.1**	-0.3	16.1**	-5.5**	-0.5†
Unemployment	1.8*	1.2**	0.1	1.6†	-0.4	0.3
Large firm	-7.1**	0.5	-0.4	-4.5**	-1.9**	-0.2
BE $> 6-12$ mths	1.2†	0.7*	-0.3†	-1.3†	1.7**	-0.6**
BE $> 12-18$ mths	1.2	-0.5	-0.2	-6.9**	3.1**	-1.3**
BE $> 18-24$ mths	-1.2	1.1	-0.3	-9.0**	2.3**	-1.4**
BE > 24 mths	-8.9**	-3.1**	-1.4**	-27.3**	0.2	-3.1**

Continued on next page...

... table 3.5 continued

ALMP measure	-28.4**	7.2**	-1.8**	-29.2**	15.2**	-1.8**
Commuter	-3.9**	-0.4	5.2**	-13.4**	0.4	8.9**
Regional characteristics						
Unemployment rate (UR)	0.1	-1.3*	0.2	-4.5*	1.0	-0.1
Change in UR 1995-2000	-1.3**	0.1	-0.1	-4.3**	0.8*	0.1
GDP per head	0.5	0.0	-0.2	-0.6	-0.3	0.2
Change of GDP 1995-2000	0.5	-0.1	0.1	0.4	0.4	-0.2
Rate of business set ups	1.3 **	-0.2	0.0	-0.3	0.2	-0.1
Share of long-term U	-3.7**	0.5†	-0.2	-3.0**	0.3	0.7**
Avg. yrs of schooling	-1.1*	-0.6*	-0.4**	0.9	-0.1	-0.2
Placement counsellor per U	-1.8**	0.2	0.3†	-2.6**	0.6†	1.4 **
Share of U in $ALMP_R$	0.5	0.1	-0.3**	0.3	0.3	-0.5**
Share of U in $ALMP_S$	0.0	0.4†	-0.1	0.6	0.8*	-0.6**
Driving time to higher level city	1.1**	0.0	-0.2†	0.6	-0.2	-0.2*
Child care places per child & child	0.8†	-0.9**	0.2†	0.9	-0.4	0.0
University present	-2.3**	0.2	0.4†	-2.2**	-0.5	0.0
Seasonal unemployment	-3.2**	3.1**	0.0	-2.5	0.4	1.5 *
Rural region	-0.4	0.2	0.4	2.9**	-1.3*	-0.4†
Urban region	1.4	0.1	-0.4*	-0.2	-0.4	-0.6**
West & 2000	13.5**	-1.1†	1.1**	6.9**	-1.6**	0.9*
West & 2001	4.4**	-1.3*	0.3	-1.9†	-0.3	0.4
East & 2000	-3.3	6.3**	1.5†	-5.9*	-0.8	8.6**
East & 2001	-0.8	3.4*	1.4†	-4.9*	-0.1	7.3**
East & 2002	2.8	3.2**	1.5*	-2.9	1.0	4.1**

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Note: Low wage sample refers to individuals with pre-unemployment daily gross wages ≤ 60 euros. BE=Benefit entitlements; ALMP=Active labour market program with focus on regular (R) or secondary (S) employment; U=Unemployment.

‡: Daily pre-unemployment wages are in the lowest (highest) wage quartile for the low (higher) wage sample.

Institutional organisation According to our findings, the recent emphasis on job counselling that is, among other things, reflected in the increasing number of job counsellors per unemployed jobseeker, is unlikely to substantially contribute to a shortening of unemployment duration. This is because a higher counsellor/customer ratio does not significantly accelerate exits from unemployment but rather affects the labour market state after unemployment. In particular, individuals with higher pre-unemployment earnings are more likely to migrate and to exit to subsidised employment while there are less local exits in regions with a higher ratio of job counsellors to unemployed jobseekers. This may suggest that additional human resources in job counselling speed up exits to migration and subsidised employment at the cost of local placement without resulting in a positive net effect on the duration of unemployment.[17] We also find that an extensive local availability of active labour market programs accelerates exits to subsidised employment at the expense of exits to regular employment. Among single persons, we mainly observe less local exits, while among married men with a higher local attachment, subsidised employment rather substitutes for non-local employment. There therefore seems to be a small regional locking-in effect of active labour market policies for married men that complements the findings from the previous chapter concerning a regional locking-in effect for women in western Germany.

Structural indicators The type of region as well as the driving time to the next large city to some extent capture differences in the availability of certain employment opportunities within a commuting range. A lack of unskilled service jobs in rural areas, for example, may help explain reduced probabilities of finding local employment for individuals with low earnings capacities. Subsidised employment partially cushions these differences with increasing exit probabilities in rural regions among low earning married men and decreasing exit probabilities for individuals with higher pre-unemployment earnings. Moreover, apart from singles with higher pre-unemployment earnings, all other groups show higher local exit probabilities in remotely located regions. This may suggest that relatively immobile groups

[17] One might retort that the counsellor/customer ratio is endogenous because the number of customers should be higher in regions with a long average unemployment duration. However, we use lagged value, i.e. the average counsellor/customer ratio during the 12 month preceding the start of the unemployment spell, which should preclude a direct simultaneity issue. Moreover, we use a number of indicators that capture regional labour market conditions such the the remaining variation in the customer/counsellor ratio is likely to mainly reflect the exogenous variation due to the politically motivated improvement of the customer/counsellor ratio during the study period.

of unemployed lower their reservation wage in regions with a lack of accessible jobs and thus experience faster exits to local employment.

The presence of a university reduces local job-finding among all groups and increases migration probabilities among all but married men with low earning capacities. These results are in line with the idea that students may exert additional congestion effects on the local labour market as students seek a minor job during their studies and often start their job search after graduation in the local area. Finally, somewhat unexpectedly, a higher level of day care places per child weakly accelerates local exits among married men with children, but not among single parents although single parents are somewhat less likely to leave a region with an extensive child care infrastructure. The relatively weak effects may partially reflect that child care facilities should only be of major importance for individuals with young children, a group that cannot be identified in the IEBS.

Western/eastern Germany Despite strong economic differences between western and eastern Germany, conditional unemployment durations are surprisingly similar. On the one hand, this suggests that our regional labour market characteristics already capture major economic differences between both parts of Germany. Since the effect of these regional characteristics are relatively limited, however, the much higher level of unemployment in eastern Germany and the long average duration of unemployment have to be explained by the huge inflow into unemployment just after reunification and the fact that many of these displaced workers never found regular employment rather than by unemployment experiences of those currently entering unemployment. For those entering unemployment between 2000 and 2002, differences between the conditional unemployment duration in eastern and western Germany are small and in many cases even disappear as we reach the end of the observation period. With regard to subsidised employment this is probably due to a reduction in the formerly extensive public spending on subsidised employment in eastern Germany. The likelihood of exiting to local regular employment, however, remains somewhat lower for most unemployed people in eastern Germany than for unemployed people in western Germany. This is partially compensated for by higher migration rates among the unemployed from eastern Germany which can be explained by the strong pull factors from western Germany.

3.7 Conclusion

In the light of recent labour market reforms, this paper explores the extent to which passive and active labour market measures, local job placement activities and regional factors affect unemployment experiences of jobseekers in Germany. For this purpose, we perform a comprehensive analysis of unemployment duration using the latest generation of administrative individual data and a broad set of regional aggregate and institutional data in the period 2000-2004. By distinguishing three exit states, local regular employment, non-local regular employment and subsidised employment we are able to disentangle the effects of individual, regional, and institutional characteristics on these destination states. This is highly relevant because there may be diverging effects on the three destination states. As a consequence, previous estimates may have been biased if non-local or subsidised employment have not been separated from exits to local employment.

Based on competing risk Cox proportional hazard estimates, we generally obtain that individual characteristics and in particular an individual's work history strongly affect the duration of unemployment and the chosen destination state while the effect of regional factors such as the unemployment rate is often rather small. This is consistent with German and international evidence concerning the impact of regional labour market conditions on the duration of unemployment until exiting to a local or non-local job (see Kettunen (2002), Yankow (2002) and chapter (2) of this dissertation). Regional disparities thus appear to be much less important than usually considered by the German public and by German policy makers. Therefore, our results suggest that regional policies may only be a supplementary means of improving labour market outcomes of unemployed individuals. Moreover, consistent with previous findings in chapter (2), we only find weak increases in migration as a response to unfavourable local labour demand conditions. Deteriorating local labour market conditions thus prolong unemployment and build up an increasing level of regional long-term unemployment.

With regard to public counselling, there is no evidence that increasing counselling efforts have much of a shortening effect on the duration of unemployment. These results may indicate that recent restructuring efforts of public employment services are unlikely to bring about a substantial reduction in unemployment. Nonetheless, restructuring efforts may contribute to the increasing efficiency of public spending, an aspect that we do not analyse

in our work. For this reason and given our econometric approach it is difficult to compare our results directly with international evaluation studies which are available for several countries, e.g. the UK (Blundell et al., 2004) and the Netherlands (van den Berg and van den Klaauw, 2006).

Entering subsidised employment in the context of an active labour market program (ALMP) often tends to cushion negative local labour market conditions and thus somewhat counteracts a prolonged unemployment duration. Moreover, previous ALMP participants are likely to end up in ALMP again, a result which likely reflects a selection of immobile individuals into what we might call ALMP-careers. Our analysis thus does not identify the causal effect of participating in ALMP on labour market outcomes which may be positive depending on the type of program as has been suggested by Lindgren and Westerlund (2003). Our analysis suggests, however, that an extensive local supply of ALMP reduces migration rates for married men. We thus find evidence for a minor regional locking-in-effect that complements the finding from chapter (2) concerning a similar effect for women.

We obtain a number of indications that the unemployment compensation and welfare system strongly affects individual labour market outcomes:

- Individuals with low pre-unemployment earnings who are likely to have high income replacement rates have the lowest exit hazards to both local an non-local regular employment.

- Exits to regular employment decrease with increasing entitlement length to unemployment benefits, especially among previously well-earning unemployed for whom exhausting unemployment benefits entails some major reductions in transfer receipt.

- Older individuals with extremely long UB entitlements basically never leave for regular employment as they use UB as a means of early retirement.

Similar to the findings from chapter (2), the evidence again suggests that the reduction of UB entitlements and income replacement rates are likely to drastically shorten unemployment and to increase migration for certain groups. A strong effect of the unemployment compensation system on the duration of unemployment has already been observed in the past. Christensen (2005) shows that social benefit recipients with high reservation wages are

unlikely to leave unemployment. Similarly, Fitzenberger and Wilke (2006b) find that unemployed people with lower former wages are much less likely to leave unemployment. Müller et al. (2007) evaluate a reform of the unemployment benefit system in 1997 which reduced entitlement length for unemployment benefits for older unemployed. They show that this reform was successful in drastically reducing inflow to unemployment and the duration of unemployment in the relevant group of unemployed.

Although our approach is fairly comprehensive and includes new data, it still has several limitations. Alternative destination states such as leaving the labour force or retirement, for example, should be an important extension to our competing risk approach. Unfortunately, our data does not provide information on these exit states such that we leave this extension to future research. Moreover, the fact that we do not observe the true length of the unemployment duration may affect our results. In addition, our econometric approach faces the methodological difficulty that a certain share of our unemployed population has zero probability for an exit to regular employment. This is known as the mover - stayer problem in the literature (Abbring, 2002; Addison and Portugal 2003) and results in the defectiveness of the unemployment duration distribution. Our estimation results may therefore be biased, but as the degree of defectiveness is limited in our data, this problem may be of minor importance. Our model does not include random effects in order to account for individual unobserved heterogeneity. For this reason we left the baseline hazard nonparametric and do not draw attention to it because it is likely to be biased. Also the assumption of proportional hazard rates can be incorrect as Fitzenberger and Wilke (2006a and 2006b) have shown with similar data that this assumption is implausible for several regressors. A more flexible approach which allows the effect of the regressors on the conditional distribution of unemployment duration to vary over the quantiles and thus even crossing of the conditional hazard rates, may provide more detailed insights. The empirical analysis in this paper still provides many new insights and it raises several interesting research questions.

Appendix

A - Summary Statistics by sub-samples

	Singles		**Married Men**	
Individual characteristics	Low Wage	Higher Wage	Low Wage	Higher Wage
Female	41.7	26.0	n/a	
Age < 26	41.4	21.2	4.4	2.5
Age 26-35	27.4	39.1	23.5	24.6
Age 46-56	10.5	12.0	30.6	28.5
Age > 56	1.8	2.0	8.0	7.0
Unskilled	52.2	30.5	45.9	32.4
University degree	1.6	7.2	1.5	5.6
Foreign born	6.2	5.3	16.9	12.4
Female foreign born	2.2	1.1	n/a	
Children	21.2	18.0	63.1	66.7
Children & female	11.6	4.8	n/a	
Minor job	8.6	1.0	8.8	1.3
Spell starts in winter	32.0	39.1	37.6	45.2
Previous employment history				
Part time	20.3	4.6	n/a	
Lower/upper wage‡	39.5	22.9	26.6	27.8
Lower/upper wage‡ & female	18.6	6.7	n/a	
Construction	4.3	2.8	6.5	4.3
Trade and Food Ind.	14.7	26.2	25.8	35.7
Services/Public sector	44.2	33.8	38.2	29.0
Recall	12.2	20.2	19.9	28.3
Unemployment	61.9	64.9	74.9	66.8
Large firm	7.0	10.3	5.1	7.0
BE $> 6 - 12$ mths	43.4	49.2	31.5	39.7
BE $> 12 - 18$ mths	1.1	2.0	3.4	4.5
BE $> 18 - 24$ mths	0.9	2.2	3.3	5.7
BE > 24 mths	0.9	2.2	3.7	7.0
ALMP measure	9.4	1.5	12.5	1.3
Commuter	20.8	27.7	20.8	29.2
Number of unemployment spells	74,724	28,168	28,018	23,620

Note: All covariates are dummy variables. BE=Benefit entitlements; ALMP= Active labour market programs; Minor Job = Job < 15 hrs/week while unemployed at the beginning of unemployment

‡: Refers to individuals with daily pre-unemployment wages in the lowest (highest) wage quartile for the low (higher) wage sample.

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Chapter 4

Bounds Analysis of Competing Risks: a nonparametric evaluation of the effect of unemployment benefits on migration in Germany

joint with Simon Lo[1] and Ralf A. Wilke[2]

Abstract

We consider a competing risks failure times model with partially identified interval data. The data problems imply that risk-specific failure distributions can only be bounded. We develop a non-parametric bounds analysis of risk-specific cumulative incidence curves (CIC) to bound a difference-in-differences effect on the CIC over different definitions of the latent durations. Our simulations demonstrate the applicability of this approach also in case of dependent competing risks. We then apply our framework to empirically evaluate the effect of unemployment benefits on observed migration probabilities in Germany. Our findings weakly indicate that reducing the maximum receipt of unemployment benefits increases the migration probability, at least among high-skilled individuals.

Keywords: cumulative incidence curve, partially missing data, difference-in-differences

JEL: C41, C14, J61

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4.1 Introduction

In this paper, we present an approach how to empirically analyse competing risks models in the case of partially missing information concerning the failure times. As an example, administrative unemployment duration data from Germany provide only incomplete information concerning the duration until leaving unemployment to different failure types because there are unobserved periods in an individual's employment trajectory (Fitzenberger and Wilke, 2004). As a result of such incomplete information, parameters of interest can only be bounded (Manski, 2003). As proposed by Abadie (2005), Lee and Wilke (2005) bound a difference-in-differences treatment effect over different definitions of the latent durations. In the context of incomplete information concerning exits from unemployment to destinations other than employment such as out of labour force or self-employment, Lee and Wilke (2005) assume independence between competing risks and thus independent censoring within a single risk framework. If this assumption holds, estimated bounds for the marginal survivor function are unbiased while they are biased in case of dependent competing risks.

The contribution of this paper is to extend the bounds framework for partially identified data to a competing risks setting and derive bounds for the empirical cumulative incidence function instead of using independent censoring as proposed by Lee and Wilke (2005). Cumulative incidence curves (CIC) refer to the empirical probability of failing from one of the causes until time t in the presence of all competing failure types (Kalbfleisch and Prentice, 1980; Pepe, 1991; Pepe and Mori, 1993). By using the CIC, our analysis yields bounds for estimates of the effects on observed failure time distributions even in case of dependent risks. It is thus a generally applicable tool to describe and analyse observed failure time distributions.

Changes of the CIC, however, convey only limited information concerning the causal effect on the underlying marginal distribution of the latent failure times. This is because the CIC does not tackle the general identification problem of competing risks models and does not recover the underlying risk-specific marginal distribution (Cox, 1962; Tsiatis, 1975). Instead, the CIC gathers information about the observed distribution of failures and thus circumvents the inherent non-identifiability of competing risks data. In light of this fundamental identification problem, nonparametric bounds of the marginal distribution as have been proposed by Peterson (1976) typically are too wide to infer some causal effect. As an

alternative, parametric assumptions can be imposed to identify the marginal distribution of a particular failure type. Under rather restrictive assumptions, Heckman and Honoré (1989) and Abbring and van den Berg (2003) show identification of the semiparametric mixed proportional hazard model. Honoré and Lleras-Muney (2006) impose quite mild assumptions to obtain tight bounds for parameters within the accelerated failure time model. Our approach avoids such parametric assumptions which are unlikely to be met. The accelerated failure time model, for example, assumes a constant effect of a regressor on the conditional risk-specific failure distributions. Since we develop our model with the purpose to apply it to a reform of the unemployment compensation system that is unlikely to have a constant effect, we require a more flexible specification. See Fitzenberger and Wilke (2007) for more details on this topic. Although our approach cannot solve the fundamental identification problem of competing risks failure models, it still provides a flexible descriptive tool for the observed distribution of competing failures under a variety of settings such as dependent risks. In particular, our approach is fully nonparametric in the sense that we do not impose assumptions that may be violated in the real world. It tackles an important identification problem that stems from missing failure time information which is present in administrative individual data if interval duration can only be partly observed. Since administrative processes often record only certain labour market states, this should be relevant for many countries. The growing literature using these data is another indicator for the relevance of our work.

After presenting the theoretical framework for a bounds analysis of risk-specific cumulative incidence curves in the case of partially missing data, a simulation demonstrates the applicability of this approach in case of dependent competing risks in contrast to a bounds analysis based on the Kaplan-Meier estimator of the marginal survivor curve. We then apply our approach to an economic research question that has already raised some research interest, namely the effect of unemployment benefits on the probability of migration. Several studies have already looked at the link between unemployment compensation and the duration of unemployment. These studies generally confirm a disincentive effect of unemployment compensation on the transitions from unemployment to employment (Katz and Meyer, 1990; Card and Levine, 2000; Lalive and Zweimüller, 2004; van Ours and Vodopivec, 2006). These findings are in line with the predictions from search theory that considers

unemployment benefits to raise reservation wages (Atkinson and Micklewright, 1991). The effect of unemployment benefits on migration, however, is much less clear. The negative effect of rising reservation wages and smaller geographical search horizons as a reaction to higher benefit levels (Hassler et al., 2005) contrasts a positive resource effect as higher unemployment benefit levels enable individuals to bear migration costs (Tatsiramos, 2003) and to increase expenditures that enhance the productiveness of job search (Barron and Mellow, 1979; Tannery, 1983). Most studies, however, seem to suggest a mobility-reducing effect of unemployment benefits on migration. Goss and Paul (1990) find that unemployment benefits reduce migration probabilities among involuntary unemployed in the US. Antolin and Bover (1997) present evidence from Spain that registered unemployed with benefit receipt are less mobile than their non-registered counterparts. Consistent with these findings, chapter (2) and (3) of this dissertation provide some evidence that unemployed in Germany who are entitled to receive unemployment benefits (UNB) for an extended period of more than 18 months are much less likely to leave unemployment and migrate than individuals with a shorter period of UNB receipt. To some extent, however, the findings of these studies may be driven by unobserved individual heterogeneity. In a study with individual fixed effects that should mitigate a possible selection problem, Tatsiramos (2003) finds a positive effect of unemployment benefits on migration, a result that he assigns to the mobility-enhancing resource effect of unemployment benefits. As a drawback, however, this study does not take account of competing transitions to local employment. Our bounds analysis thus reexamines the effect of shorter unemployment benefit receipt on the transitions to either local or non-local employment via migration. For this purpose, we exploit a natural experiment that generates some exogenous variation of entitlement length, namely the reform of unemployment benefit entitlements in Germany in 1997. This reform reduced the length of entitlements for certain age groups by up to ten months. As a consequence, it is possible to construct a treatment group with shortened entitlement length and a control group for whom entitlements have been unaffected by the reform. Based on two different definitions of the latent durations until exiting to either local or non-local employment, it is possible to apply our bounds framework to the inference of the effect on the CIC for migration and for other exit states. The findings indicate that the degree of uncertainty in the German administrative data that is due to unobserved periods in an individual's employment tra-

jectory at first precludes any clear inference as the bounds tend to be very wide. When introducing an additional assumption concerning the missing information, bounds are generally much tighter. For high-skilled individuals, for whom the threat of entitlement loss due to the 1997 reform is likely to be largest, the corresponding bounds are indicative for the mobility-reducing effect of extensive UB receipt.

The paper is structured as follows. The following section presents the theory of a competing risks bounds analysis of cumulative incidence curves in the case of partially missing information. We then simulate the applicability of this approach in case of non-random censoring and dependent competing risks. Section 4.4 applies the proposed method to analyse the effect of unemployment benefits on the cumulative incidence of migration. Section 4.5 concludes.

4.2 Model

Let the pair (T, R) be the latent failure times and exit states, respectively, where the failure times are observed on a discrete scale $t_1 < t_2 < \ldots < t_k$ and exit states $R = 1, \ldots, r, \ldots, z$. Latent censoring time is denoted by C. There are $i = 1, \ldots, N$ independent and identically distributed observations. Failure time for risk r of observation i, $(T_i, r) = t_j$ is observed only if $(T_i, r) = min_R\{(T_i, R), C_i\}$. We assume independent censoring throughout the paper, i.e.

$$\begin{aligned} \lambda_{rj} &= P[(T,r) = t_j | (T,1) \geq t_j \cap \ldots \cap (T,r) \cap \ldots \cap (T,z) \geq t_j \cap C \geq t_j] \\ &= P[(T,r) = t_j | (T,1) \geq t_j \cap \ldots \cap (T,r) \cap \ldots \cap (T,z) \geq t_j], \end{aligned} \tag{4.1}$$

which means that the censoring in the data does not change the risk-specific distribution of failure times and thus the hazard rate λ_{rj}. We suppose further that failure type r and censoring occurs with multiplicity d_{rj} and c_j at time t_j, respectively. If all competing risks are independent, the hazard rate from (4.1) becomes

$$\lambda_{rj} = P[(T,r) = t_j | (T,r) \geq t_j], \tag{4.2}$$

and the likelihood function can be written as

$$L = \prod_{j=1}^{k} \left\{ \prod_{r=1}^{z} P[(T,r) = t_j]^{d_{rj}} P[C = t_j]^{c_j} \right\}. \tag{4.3}$$

The likelihood L can be maximised by replacing $P[C = t_j]$ by $P[T > t_j] = P[(T,1) > t_j \cap \ldots \cap (T,z) > t_j]$ as the only contribution to the likelihood by the censored observations. After rearranging,

$$L = \prod_{j=1}^{k} \left\{ \prod_{r=1}^{z} \lambda_{rj}^{d_{rj}} (1-\lambda_j)^{n_j - d_j} \right\}, \tag{4.4}$$

where $\lambda_j = \sum \lambda_{rj}$, $d_j = \sum d_{rj}$ and $n_j = (d_j + c_j) + \ldots + (d_k + c_k)$ represents the number of observations at risk. Using the first-order-condition, the Kaplan-Meier estimate of λ_{rj} which maximises (4.4) and thus the (overall) survivor function, $\hat{S}(t_j)$, and the risk-specific cumulative incidence curve (CIC), $\hat{I}_r(t_j)$, are given by

$$\hat{\lambda}_{rj} = \frac{d_{rj}}{n_j}, \quad \text{and} \tag{4.5}$$

$$\hat{S}(t_j) = \hat{P}[T > t_j] = \prod_{l=1}^{j} (1 - \hat{\lambda}_l) = \prod_{l=1}^{j} \frac{n_l - d_l}{n_l}, \tag{4.6}$$

$$\hat{I}_r(t_j) = \hat{P}[(T,r) \leq t_j] = \sum_{l=1}^{j} \hat{\lambda}_{rl} \hat{S}(t_l). \tag{4.7}$$

It can be shown that, in the independent competing risks setting, (4.5) has the same value if we treat the competing risks other than r as censored, i.e. replacing $P[(T, R \neq r) = t_j]$ by $P[T > t_j]$ in (4.3). In other words, censoring is equivalent to independent competing risks with unknown exit states. For the survivor function, however, the definition of competing risks and censoring are not interchangeable. Since λ_j is a sum of all risk-specific hazard rates, treating some of the competing risks as censored changes the estimated (overall) survivor curve as in (4.6) to reflect only the probability of survival from the remaining competing risks. In the case of independent competing risks, treating all competing risks besides r as censored still produces consistent results. The Kaplan-Meier survivor curve from (4.6) becomes a consistent estimator for the marginal survivor curve specific to risk r only, i.e. (4.4) and (4.6) becomes

$$L^c = \prod_{j=1}^{k} \left\{ \lambda_{rj}^{d_{rj}} (1-\lambda_{rj})^{n_j - d_{rj}} \right\}, \text{ and} \tag{4.8}$$

$$\hat{S}_r^c(t_j) = \prod_{l=1}^{j} (1 - \hat{\lambda}_{rl}) = P[(T,r) > t_j]. \tag{4.9}$$

In the case of dependent competing risks, the above Kaplan-Meier type estimator is biased. See Moeschberger and Klein (1995) for a review. As hazard rates are now interdependent,

competing risks cannot arbitrarily be considered as censored observations. Nevertheless, as is shown by Cox (1962) and Tsiatis (1975), from any observed distribution $(\tilde{T}, r)$ with dependent joint probability $P[(\tilde{T}, R) > t_j] = P[(T, 1) > t_j \cap \ldots \cap (T, z) > t_j]$ an independent joint probability function $(\acute{T}, R)$ can always be fitted which produces the same observed distribution, i.e.

$$\begin{aligned} P[(\tilde{T}, R) > t_j] &= P[(T, 1) > t_j \cap \ldots \cap (T, z) > t_j] \\ &= P[(\acute{T}, 1) > t_j] \ldots P[(\acute{T}, z) > t_j]. \end{aligned} \tag{4.10}$$

$P[(\acute{T}, r) = t_j]$ is the marginal density function of the hypothetical distribution of value zero if r is not observed at t_j, and thus $P[(\tilde{T}, r) = t_j] = P[(\acute{T}, r) = t_j]$ holds for all r and t_j. Thus, the Kaplan-Meier estimator from (4.6) still produces unbiased estimates of the overall survivor curve when replacing the true failure time T with the assumed independent $\acute{T}$ in (4.2) to (4.6) and treating all dependent risks as competing risks. It is, however, no longer a consistent estimator for the marginal survivor curve specific to risk r. This is the result of the non-identification problem, and thus the underlying marginal probability for each failure cannot be identified without any additional parametric assumptions.

The cumulative incidence curve (CIC) has been suggested as an alternative nonparametric tool which has a meaningful interpretation also in presence of dependent competing risks (Kalbfleisch and Prentice, 1980; Pepe, 1991; Pepe and Mori, 1993). The CIC refers to the observed probability of experiencing a specific failure type prior to time t_j in the presence of all competing failure types. It therefore does not recover the underlying risk-specific marginal distribution. Instead, it refers to observed failure probabilities which are also well defined in the case of dependent competing risks. In other words, the CIC offers a descriptive tool.

In what follows, we present bounds for the CIC in a context of partially identified data. In such a setting, (T_i, z) is not exactly identified and we assume a lower- and upper-bounded latent distribution for z, i.e. $(T^{LB}, z) < (T, z) < (T^{UB}, z)$. The latent distribution for the remaining risks, $(T, R \neq z)$, are fully identified and have no bounds, i.e.

$$\begin{aligned} (T^{LB}, R) &= \{(T, 1), \ldots, (T, z-1), (T^{LB}, z)\}, \text{ and} \\ (T^{UB}, R) &= \{(T, 1), \ldots, (T, z-1), (T^{UB}, z)\}. \end{aligned}$$

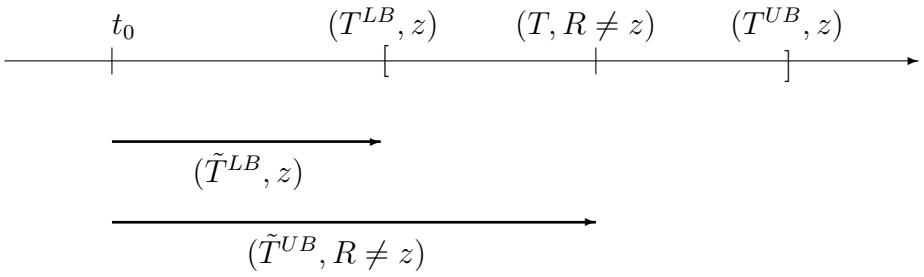
Thus, we also observe a lower bound and upper bound for each risk-specific duration time, $(\tilde{T}_i^{LB}, R)$ and $(\tilde{T}_i^{UB}, R)$. This is because the observed distribution for risks $R \neq z$ depends

on the bound for risk z and the bounded failure data on risk z will change the observations on risks other than z. If there was no missing information problem, the observable duration time for risk z is $(\tilde{T}_i, z)$. Instead, given the partial identification problem, if we observe a failure to risk z, then we have for sure observed the lower bound $(\tilde{T}_i^{LB}, z)$. The upper bound for this duration is defined by $min\{(T_i, r), (T_i^{UB}, z), C_i\}$ for $r = 1, \ldots, z-1$. This means that the upper bound may differ not only in terms of duration but also in terms of the exit state:

$$\begin{aligned}
(i) \qquad & (\tilde{T}_i^{LB}, z) \leq (\tilde{T}_i, z) \leq (\tilde{T}_i^{UB}, z), \text{ or} \\
(ii) \qquad & (\tilde{T}_i^{LB}, z) \leq (\tilde{T}_i, z) \leq (\tilde{T}_i^{UB}, R \neq z), \text{ or} \\
(iii) \qquad & (\tilde{T}_i^{LB}, z) \leq (\tilde{T}_i, z) \leq C_i^{UB}
\end{aligned}$$

As an illustration, Figure 4.1 shows the case where the true $(\tilde{T}_i, z)$ falls in region (ii). Under the definition of the lower bound, risk z is observed at $\tilde{T}^{LB}$, whereas risk $R \neq z$ is observed at $\tilde{T}^{UB}$ under the definition of the upper bound.

Figure 4.1: Upper and lower bound of the observed risk specific duration



Regarding the empirical section, the particular data problem is that only a lower bound for exit state z is observed, and thus only regions (ii) and (iii) are valid. Formally speaking, the model studies the latent distributions of $(T, 1), \ldots, (T, z-1)$ and $min\{T_{R \neq z}, T_z\}$, as (T, z) does not necessarily fall into the bounds indicated by the second and third region if it is not the minimum of (T, R) and thus could never be observed.

The hazard rate, survivor curve and the CIC are estimated nonparametrically using the above equations by replacing $\tilde{T}$ with $\tilde{T}^{LB}$ and $\tilde{T}^{UB}$ respectively. In analogy to Lee and Wilke (2005), we use the monotone relations of the survival function and the CIC to formulate the bounds analysis and to study the treatment effect of some policy reforms. Using the

relations $\tilde{T}^{LB} \leq \tilde{T} \leq \tilde{T}^{UB}$, it is straightforward to see the following relations, which hold for $r = 1, \ldots, z-1$:

$$\begin{aligned} d_{rj}^{LB} &\leq d_{rj}^{UB}, \\ \sum_{l=1}^{j} d_{zl}^{LB} &\geq \sum_{l=1}^{j} d_{zl}^{UB}, \\ n_j^{LB} &\leq n_j^{UB}, \text{ and thus} \end{aligned}$$

$$\hat{S}^{LB}(t_j) \leq \hat{S}(t_j) \leq \hat{S}^{UB}(t_j) \tag{4.11}$$

$$\hat{I}_r^{LB}(t_j) \leq \hat{I}_r(t_j) \leq \hat{I}_r^{UB}(t_j) \tag{4.12}$$

Bounds for other functions such as the cause-specific hazard rate or the cause-specific cumulative hazard rate cannot be derived. Thus, the identification problem concerning the latent duration (T, z) implies that we observe only bounds for all risk-specific survivor curves and CICs. The implications of dependent competing risks for these bounds call for some remarks.

Remark 1 If the exit state z is independent to the remaining competing risks and we treat it as censored, the distributions of the fully observed failure times are then estimated in a way without making use of the additional information provided by this risk, and thus the bounds analysis is obsolete[3], i.e.

$$\hat{S}_r^{c,\,LB}(t_j) = \hat{S}_r^{c,\,UB}(t_j) = P[(T,1) > t_j \cap \ldots \cap (T, z-1) > t_j], \text{ and} \tag{4.13}$$

$$\hat{I}_r^{c,\,LB}(t_j) = \hat{I}_r^{c,\,UB}(t_j) = \sum_{l=1}^{j} P[(T,r) = t_l]. \tag{4.14}$$

(4.13) is the unbiased estimator for the survivor curve of the remaining competing risks and - since risk z is independent - of the overall survivor curve. (4.14) is the latent marginal density of exit r.

Remark 2 If risk z is dependent, the previously discussed properties of the survivor curve and the CIC under competing risks carry over to their respective bounds. Moreover, the estimated bounds depend on the chosen upper and lower bound for the unobserved latent

[3]This follows from the definition of independent censoring. Kaplan-Meier type estimators are consistent estimators for the survivor curve and the CIC of the latent marginal distribution. Since this property has an asymptotic nature, there may be some deviations in an application. We may observe a slight discrepancy of the lower and upper bound for which a monotone relation does not necessarily hold.

durations. If we treat risk z as censored, the bounds of the survivor curve bound the survivor curve for the hypothetical independent distribution and not the actual overall survival function. The bounds for the CIC exist even if risk z is dependent and we treat it as a competing risk because $\acute{I}_r^{LB}(t_j) = \sum_{l=1}^{j} P[(\tilde{T}^{LB}, r) = t_l]$ and $\acute{I}_r^{UB}(t_j) = \sum_{l=1}^{j} P[(\tilde{T}^{UB}, r) = t_l]$ and the monotone relation $P[(\tilde{T}^{LB}, r) < t_j] > P[(\tilde{T}, r) > t_j] < P[(\tilde{T}^{UB}, r) > t_j]$ implies $\acute{I}_r^{LB}(t_j) < I_r(t_j) > \acute{I}_r^{UB}(t_j)$. If the dependent risk z is treated as censored, bounds of the hypothetical risk-specific CIC will not have a meaningful interpretation because the information provided by the dependent risk z is dropped artificially.

To conclude the above argument, bounds analyses for the CIC and the survivor curve can generally be used by treating missing data as a competing risk instead of as censored. By doing so, we estimate the unbiased overall survivor curve and the risk-specific CIC.

Now consider a setting where the duration of interest is the unemployment duration. Moreover, an individual faces several risks: it may enter local employment, non-local employment, or it may leave the labour force, become self-employed or enter subsidised employment. Unfortunately, the data offers only limited information on certain exit states $R = 1, \ldots, z$, i.e. $(\tilde{T}_i, R)_{R=1,\ldots,z}$ can be observed while all other competing risks cannot be distinguished. Thus, for the indistinguishable other exit states denoted with $R = o$, we observe $(\tilde{T}_i, o) = min\{(T_i, o_1), (T_i, o_2), \ldots, (T_i, o_v)\}$, with (T, o_v) representing the latent failure times for the non-distinguishable exit states o_v.[4] Lower and upper bounds are denoted with $(\tilde{T}_i^{LB}, o)$ and $(\tilde{T}_i^{UB}, o)$. Duration is independently censored with time C_i. The duration to exit state o is dependent on that of exit state $R = 1, \ldots, z$. Treating the exit state o as a competing risk, bounds for a treatment effect on the overall survivor curve and on cause-specific cumulative incidence curves can be used without imposing the independence assumption.

Now, suppose there is a policy intervention in a natural experiment setting. We have two groups, the control group ($G = 0$) and the treatment group ($G = 1$), and two time intervals, the pre-reform period ($P = p_{t0}$) and the post-reform period ($P = p_{t1}$). The reform of interest is supposed to have an effect on the unemployment duration of the treatment group in the post-reform years and the effect of the reform can be estimated by a Difference-in-Differences

[4] In an application, the pooling of all unidentified exit states aggravates its interpretability as changes in the failure time for certain exit states subsumed under the exit state o may actually oppose each other.

estimator (DID) as

$$\Delta_{Ir}(t_j|p_{t0},p_{t1},x) = [I_r(t_j|1,p_{t1},x) - I_r(t_j|0,p_{t1},x)] - [I_r(t_j|1,p_{t0},x) - I_r(t_j|0,p_{t0},x)] \quad (4.15)$$

for $r = 1,\ldots,z$, where $I_r(t_j|g,p,x) = P(T \leq t_j, R = r, G = g, P = p, X = x)$ with X as further observable variables such as gender, age etc. and

$$\Delta_S(t_j|p_{t0},p_{t1}) = [S(t_j|1,p_{t1},x) - S(t_j|0,p_{t1},x)] - [S(t_j|1,p_{t0},x) - S(t_j|0,p_{t0},x)], \quad (4.16)$$

where $S(t_j|g,p,x) = P(T > t_j, G = g, P = p, X = x)$. Given that we observe bounds for the estimated survivor function as in (4.11) and bounds for the estimated cause-specific cumulative incidence curve as in (4.12) it is straightforward to derive bounds for $\hat{\Delta}_{Ir}$ and $\hat{\Delta}_S$ since they are bounded by intervals with endpoints $[l_{Ir}(t_j|p_{t0},p_{t1},x), u_{Ir}(t_j|p_{t0},p_{t1},x)]$:

$$\begin{aligned} l_{Ir}(t_j|p_{t0},p_{t1},x) = \max[-1, &\{I_r^{LB}(t_j|1,p_{t1},x) - I_r^{UB}(t_j|0,p_{t1},x)\} \\ &- \{I_r^{UB}(t_j|1,p_{t0},x) - I_r^{LB}(t_j|0,p_{t0},x)\}] \end{aligned} \quad (4.17)$$

and

$$\begin{aligned} u_{Ir}(t_j|p_{t0},p_{t1},x) = \min[1, &\{I_r^{UB}(t_j|1,p_{t1},x) - I_r^{LB}(t_j|0,p_{t1},x)\} \\ &- \{I_r^{LB}(t_j|1,p_{t0},x) - I_r^{UB}(t_j|0,p_{t0},x)\}] \end{aligned} \quad (4.18)$$

for $r = 1,\ldots,z$ and $[l_S(t_j|p_{t0},p_{t1},x), u_S(t_j|p_{t0},p_{t1},x)]$:

$$\begin{aligned} l_S(t_j|p_{t0},p_{t1},x) = \max[-1, &\{S^{LB}(t_j|1,p_{t1},x) - S^{UB}(t_j|0,p_{t1},x)\} \\ &- \{S^{UB}(t_j|1,p_{t0},x) - S^{LB}(t_j|0,p_{t0},x)\}] \end{aligned} \quad (4.19)$$

and

$$\begin{aligned} u_S(t_j|p_{t0},p_{t1},x) = \min[1, &\{S^{UB}(t_j|1,p_{t1},x) - S^{LB}(t_j|0,p_{t1},x)\} \\ &- \{S^{LB}(t_j|1,p_{t0},x) - S^{UB}(t_j|0,p_{t0},x)\}]. \end{aligned} \quad (4.20)$$

Note that the lower and upper bounds are restricted to be between -1 and 1. This is due to the fact that maximum variation of probabilities cannot be larger than 1 in absolute values. In the next section, we use simulations to demonstrate our major findings in this section.

4.3 Simulations

The following simulation framework is closely related to the empirical application which is concerned with transitions from unemployment. For the simulation, we thus assume that there are two observed exit states, employment ($r = e$) and other destinations ($r = o$) such as leaving the labour force and self-employment and no censoring.[5] $\tilde{T}_o$ is not directly observed. Let $\tilde{T}^{LB}$ be the observed lower bound of $\tilde{T}_o$. To simplify the issue, we use $\tilde{T}_e$ as a natural upper bound of $\tilde{T}_o$.[6] We assume that the latent failure times, T_e and T^{LB}, have the following multivariate lognormal distribution

$$log(T_r) \sim N(\mu_r + G \times c_{1r} + D_P \times c_{2r} + G \times D_P \times c_{3r}, \Omega),$$

$$\Omega = \begin{pmatrix} \sigma_e^2 & \rho_{eLB}\sigma_e\sigma_{LB} \\ \rho_{eLB}\sigma_e\sigma_{LB} & \sigma_{LB}^2 \end{pmatrix},$$

with $r = e, LB$. G is a group dummy equal to zero for the control group and equal to one for the treatment group. D_P is a period dummy equal to zero in the pre-reform period and equal to one in the post-reform period. μ thus refers to the mean failure time for the control group in the pre-reform period. Values of the parameters $\{\mu_r, c_{ir}, \sigma_r\}$ for r=e, LB and i=$1, 2, 3$ are chosen to approximate the conditions in the empirical application and are summarised in Table 4.1:

Table 4.1: Chosen parameters for the simulation framework

r	μ	$c1$	$c2$	$c3$	σ
e	1.0	0.4	0.6	-0.4	2.5
LB	0.8	0.0	0.0	-0.5	1.5

[5]In the application, we will further distinguish between employment in the local area and employment via migration.

[6]If $\tilde{T}_o$ was larger than $\tilde{T}_e$, it could not be observed and therefore did not exist.

We further assume that the true failure time for exit o, T_o, follows a uniform distribution in the interval $[T_{LB}, T_{LB} + 5]$. Moreover, the treatment group is assumed to have longer latent failure times to all exit states in both the pre- and post-reform period than the control group ($c_1 > 0$). Due to factors unrelated to the reform, both the treatment and control group have longer failure times to all exit states in the post-reform period than in the pre-reform period ($c_2 > 0$). The interaction term $G \times D_P$ reflects the assumption that the reform decreases the failure time to exit state e on the treatment group only ($c_3 < 0$). We first adopt a framework with zero correlation between the failure time to exit state e and its lower limit, i.e. ρ_{eLB}=0, in order to study the case of independent competing risks before we relax this assumption later.

Using formulas (4.15)-(4.16), we make use of the fully identified competing risks, T_o, to compute the estimated change in the observed distribution for exit state e. We then compute the bounds for these changes for the situation that we observe only T^{LB}. We apply the formulas (4.17)-(4.20) and treat T^{LB} either as censored or as a competing risk. We run 500 simulations, each with $10,000$ failure times for each exit state to derive consistent estimates for the changes of the observed distribution functions and the corresponding bounds. The mean value and the bounds for the effects are reported in Figure 4.2. In addition, the 5% (95%)-quintile of the estimated lower bounds (upper bounds) are also included.

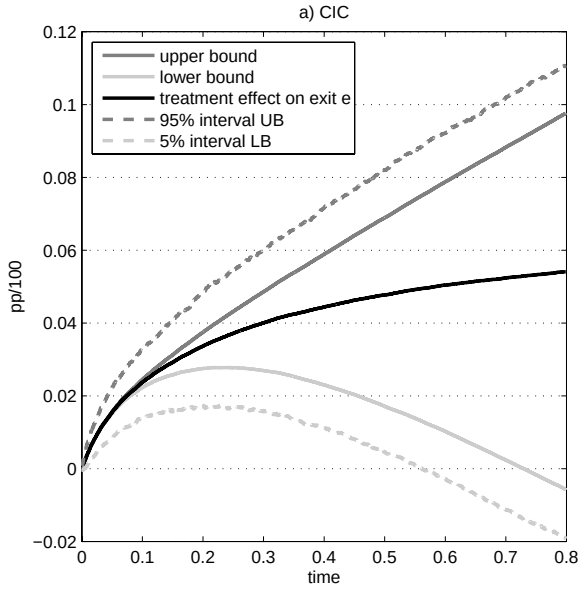
As the parameters c_{3e} and c_{3o} are negative, failure times to all exit states are shortened after the reform. The expected positive treatment effect on the CIC for exit e and the expected negative effect on the overall survivor curves with fully identified data are shown as the black thick line in Figure 4.2(i)(a), (ii)(a) and Figures 4.2(i)(b), (ii)(b), respectively.[7]

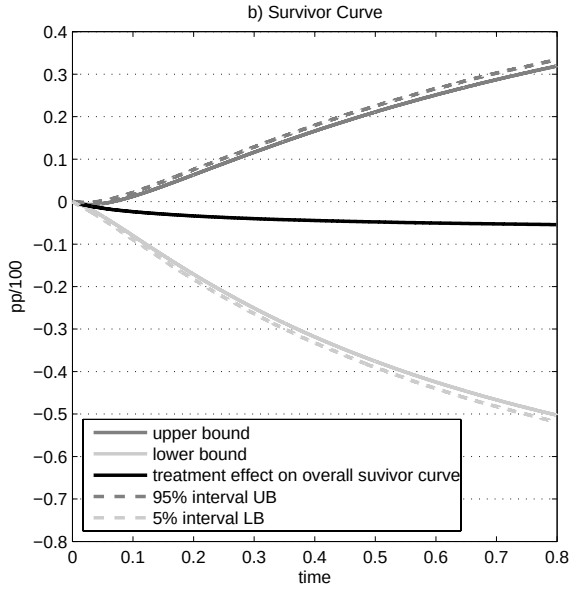
In Figure 4.2(i)(a) & (b), results using the competing risks framework show that the bounds contain the estimated treatment effect on the CIC for exit e and the overall survivor curve. Using the censoring framework, as discussed in remark 1 in the last section, the upper and lower bounds coincide with each other in Figure 4.2(ii)(a) & (b) and are an unbiased estimate for the treatment effect on the latent marginal survivor curve of exit e. As T_{LB} and T_o are independent of T_e, the overall survivor curve and the CIC in the presence of T_o correspond to the estimated marginal survivor curve.

[7]In both the competing risk and the censoring setting, the estimated "true" reform effects are computed in the same way.

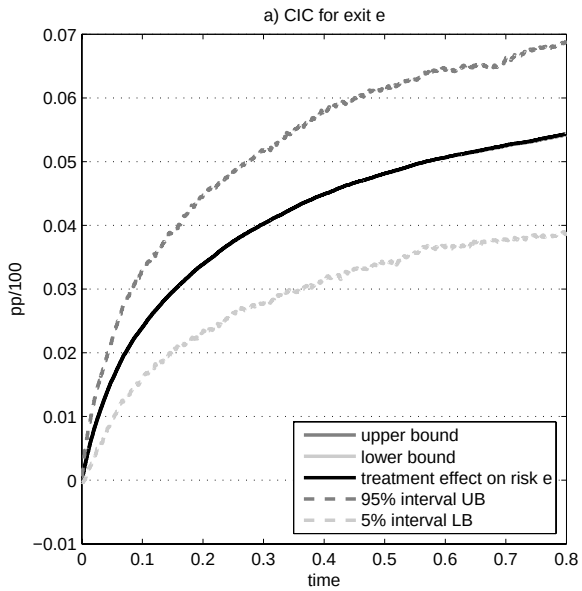
Figure 4.2: Simulated treatment effects and their bounds in case of uncorrelated failures types

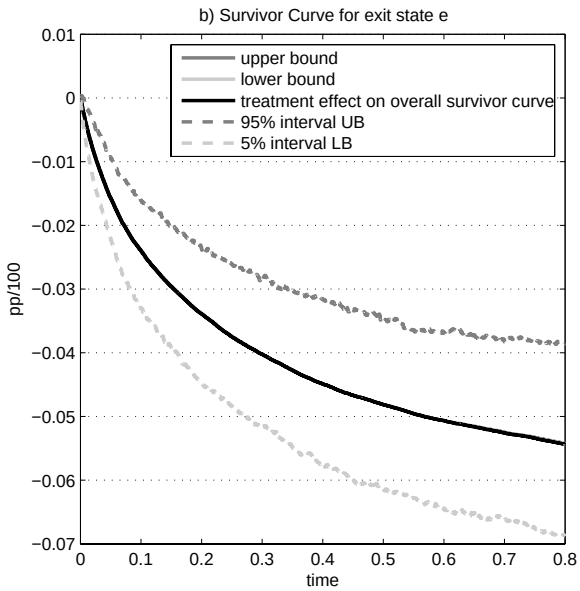
(i) Treating T^{LB} as a competing risk





(ii) Treating T^{LB} as censored



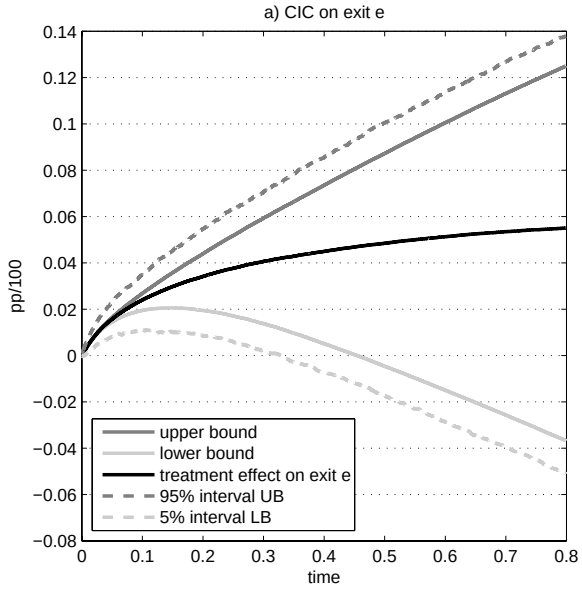


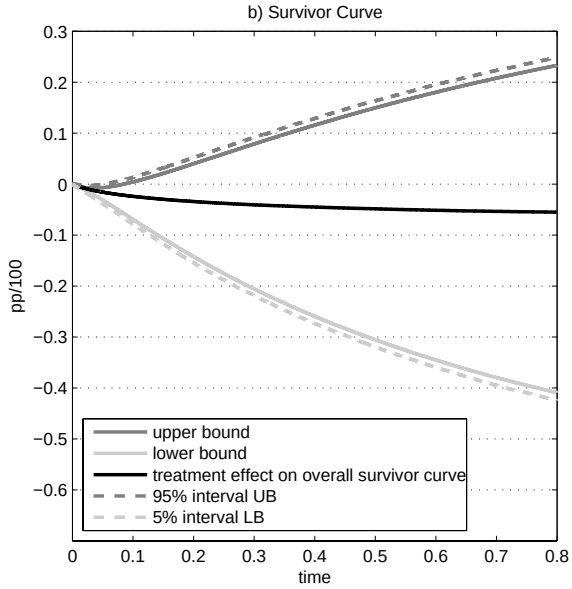
Next we allow a correlation of T_e and T^{LB} with ρ_{eLB}=0.05 and thus introduce a setting in which the failure times for the exit state e and the observed lower bound are dependent. In a competing risk framework, Figure 4.3(i) is very similar to Figure 4.2(i). The bounds contain the estimated treatment effect on both the overall survivor and the cause-specific cumulative incidence curve. In the presence of dependent competing risks, the bounds for the marginal survivor curve of exit e revert in direction when treating T^{LB} as censored as shown in Figure 4.3(ii)(b) and are thus not applicable. In fact, as discussed in remark 2 of the last section, the bounds are now bounding neither the treatment effect on the overall survivor curve nor the marginal survivor curve for the employment. When treating T^{LB} as censored, Figure 4.3(ii)(a) shows that a slight increase in the correlation ρ_{eLB} from zero to 0.05 suffices to dramatically widen the bounds from a single line to a wide bound. This bound differs from that without censoring in 4.3(i)(a) and, as discussed in remark 2 of the last section, does not have a meaningful interpretation.

We conclude that estimating the bounds on the CIC and treating T^{LB} as a competing risk is a preferable approach. In the case of independent competing risks, both bounds on the CIC and the survivor curve produce a consistent estimate of the latent marginal treatment effect. In the case of dependent competing risks, however, the bounds for the survivor curve are now bounding neither the treatment effect on the overall survivor curve nor the marginal survivor curve for the exit state employment. By contrast, the bounds for the CIC are still interpretable as an estimate of the effect on the observed distribution of failures, but - due to the non-identifiability of competing risks - do not any longer capture the latent marginal treatment effect. Since the degree of dependence between competing risks is unknown, a bounds analysis of CICs is a preferable approach as it yields an interpretable result under all conceivable circumstances.

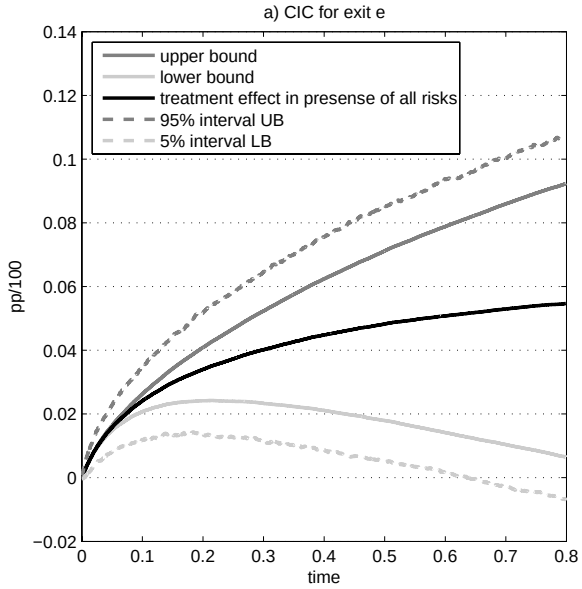
Figure 4.3: Simulated treatment effects and their bounds in case of ρ_{el}= 0.05

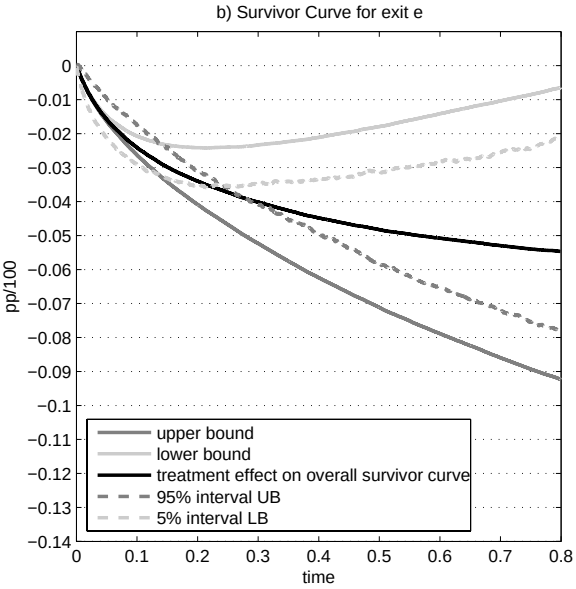
(i) Treating T^{LB} as a competing risk





(ii) Treating T^{LB} as censored





4.4 Empirical Application

As briefly discussed in the introduction, we apply our bounds analysis to bound the effect of reducing the maximum duration of receiving unemployment benefits on the observed transitions from unemployment to local and non-local employment via migration. We begin this section with a brief description of the German unemployment compensation system and discuss the 1997 reform of unemployment benefit entitlements. This discussion is based on the Employment Promotion Act (*Arbeitsförderungsgesetz*), the Social Welfare Act III (*Sozialgesetzbuch III*) and several secondary sources such as Plaßmann (2002), Oschmiansky et al. (2001) and Wolff (2003). We then introduce the data and discuss the selection of treatment and control group, before we present the result of bounding the effect as described in the previous methodological section.

Basic features of the unemployment compensation system During the study period, the system of unemployment compensation in Germany consists of two main components: unemployment benefits (UNB) and unemployment assistance (UNA). As an insurance, unemployment benefits are limited in time depending on the length of socially insured employment during a period of seven years before the benefit claim. Moreover, the length of benefit receipt positively depends on age with a maximum UNB receipt of 12 months for younger age groups and up to 32 months for older age groups in the years prior to the 1997 reform. After exhausting UNB, unemployed individuals receive the tax-funded unemployment assistance if they pass a means-test. Both UNB and UNA are a percentage of former wage income with UNB replacing 63% (68%) of former wage income and UNA still reaching income replacement rates of 53% (57%) for individuals without (with) dependent children. For individuals with low pre-unemployment wages, income replacement rates may even be higher. If the unemployment compensation as a percentage of former wage income does not suffice to ensure the legally defined minimum standard of living, individuals receive complementary social benefits. As a result, income replacement rates for individuals with low pre-unemployment wages may be close to 100% and disincentives to take up a new job should be particularly severe for this group of unemployed. Consistent with such disincentive effects, there is empirical evidence that these groups experience longer unemployment durations and are less likely to leave for non-local jobs than unemployed individuals with

higher pre-unemployment wages (see chapter (3) of this dissertation). The design of the unemployment compensation and welfare system in Germany thus implies that any changes concerning the maximum duration of UNB receipt are ineffective for unemployed individuals with complementary social benefits. Receiving unemployment assistance instead of unemployment benefits does not change the income replacement rate for these individuals and should thus not affect their job search strategy. By contrast, unemployed individuals without additional social benefits but with eligibility for the means-tested UNA loose around 10% of their former wage income when switching from UNB to UNA. For this group, a shortening of UNB is likely to have a small effect only. Individuals who do not pass the means test for receiving UNA due to other income sources or private savings even loose all unemployment compensation after exhausting UNB. The threat of entitlement loss should thus be strongest for this rather small group of unemployed.

1997 Reform In April 1997, a major reform of the Employment Promotion Act came into force to shorten the receipt of UNB for some of the older age groups and to introduce stricter sanction rules for the non-compliance with certain eligibility requirements. The enforcement of stricter sanction rules in Germany after 1997 may have accelerated the transition from unemployment to employment because temporary reductions in UNB due to non-compliance with eligibility rules have been found an effective means of reducing unemployment (Boone et al., 2002, 2004). Since these new regulations applied to all unemployed at the same time, however, our DID framework should eliminate this effect and still allow for the identification of the causal effect of shortening the UNB receipt for some older age groups in 1997. In Germany, the potential UNB duration (PUNBD), i.e. the maximum duration of UNB receipt at the beginning of the unemployment period, positively depends on the period of socially insured employment within the seven years prior to the benefit claim. This so called extended claim period is restricted by previous benefit claims and is thus shorter than seven years for individuals with a benefit claim within the previous seven years. In addition, the PUNBD positively depends on age. During the 1980s, the PUNBD had successively been expanded for older age groups. Thus, before the reform in 1997, entitlements to UNB lasted up to 32 months for individuals above the age of 42, while the PUNBD for individuals below this age range was only 12 months. A detailed description of these earlier reforms can be found in Hunt (1995). One well-documented result of these earlier reforms that demonstrates the

disincentive effect of this system was the rapid increase of early retirees whose extremely long UNB receipt allowed for bridging the gap between employment and retirement age. See Fitzenberger and Wilke (2004) for a nonparametric analysis using similar administrative data. In 1997, the PUNBD was reduced for some of the older age groups by lowering the age limits for certain maximum UNB receipts (see Table 4.2). As a consequence, the PUNBD for individuals between 42 and 43 years of age was cut from 18 months before 1997 to 12 months after the 1997 reform. For individuals aged 44, UNB was even cut from a maximum receipt of 22 to a maximum receipt of 12 months. Individuals aged below 42 years were unaffected by the reform as they always received a maximum of 12 months of UNB. The 1997 reform thus provides a natural experiment with a credible source of variation in PUNBD that can be used to identify its causal effect.

Table 4.2: Potential unemployment benefit duration (PUNBD) for UNB claimants up to age 47 by work history and age, IAB-R01

Soc. insured employment	**PUNBD (in months)**	
during claim period	until 03/97	since 04/97
12 months	6	6
16 months	8	8
20 months	10	10
24 months	12	12
28 months	14 (age $\geq$42)	14 (age $\geq$45)
32 months	16 (age $\geq$42)	16 (age $\geq$45)
36 months	18 (age $\geq$42)	18 (age $\geq$45)
40 months	20 (age $\geq$44)	20 (age $\geq$47)
44 months	22 (age $\geq$44)	22 (age $\geq$47)

Source: Plaßmann (2002)

One problem of the 1997 reform that has to be taken into account, however, is that the implementation of the reform was partially cushioned. Until March 1999, new benefit claimants were treated according to the pre-reform regulations if there was a work history of more than one year during the three years prior to the benefit claim. Thus, the new regulations applied to all new benefit claims after March 1999 only. Two German studies already looked at the effect of the 1997 reform on transitions from unemployment to employment. Based on the German Socio-Economic Panel, Wolff (2003) only finds very weak positive

effects of shortening the PUNBD on the transitions to employment in eastern Germany. As previously discussed, this finding may reflect that the entitlement loss due to the 1997 reform was rather limited for most groups. Moreover, due to the limited sample size of the GSOEP data, the study pools unemployment spells starting between 1990 and 1999 and thus includes only a limited number of spells that were actually affected by the reform. In the subsequent analysis, we use an administrative data set that provides a much larger sample size and thus also allows for distinguishing between exits to local versus exits to non-local employment after migration. Based on the same data set, Müller et al. (2007) look at the effect of the 1997 reform on transitions to employment among older unemployed above the age of 52 for whom the 1997 reform also shortened the PUNBD. They find evidence that the reform reduced the inflow into unemployment and drastically reduced the duration of unemployment among this group, a result that suggests that shorter UNB durations lower the attractiveness of early retirement via the receipt of UNB. Using the same administrative data set, we reexamine the effect of the PUNBD on transitions to local and non-local employment. Moreover, we restrict the analysis to prime age individuals for whom early retirement should not be an issue. Moreover, prime age individuals are much more likely to migrate as a response to the 1997 reform than their older counterparts.

Data: IAB-R01 The analysis is based on the IAB employment subsample 1975-2001 - regional file (IAB-R01[8]). This register data set contains spell information on a 2 % sample of the population working in jobs that are subject to social insurance payments and thus excludes self-employed individuals and tenured civil servants. The data contains spell information on periods for which the individual received unemployment compensation (UC) from the Federal Employment Agency (*Bundesagentur für Arbeit*) such as unemployment benefits UNB (*Arbeitslosengeld*), unemployment assistance UNA (*Arbeitslosenhilfe*) and maintenance payments during training measures MP (*Unterhaltsgeld*). Thus, employment histories including periods of transfer receipt can be reconstructed on a daily basis. One major drawback of the data set is that the true unemployment duration is not known because the data only contains information on the receipt of UC. As a consequence, there is a gap in the IABS-R01 record whenever an individual continues to be unemployed after exhausting unemployment

[8]See Hamann et al. (2004) for a detailed description of the IAB-R01.

benefits without receiving unemployment assistance. Since such a gap in the IABS-R01 record is indistinguishable from other unobserved labour market states, such as being out of the labour force or self-employed, there is uncertainty about the true duration until leaving unemployment to one of these other destination states (o). As a consequence of this partially missing information problem, it is necessary to define unemployment spells according to a suitable bound (Fitzenberger and Wilke, 2004, Lee and Wilke, 2005). For the following analysis, we use two proxies, an upper and a lower bound that can be used for the bounds analysis as discussed in the methodological section. In our case $\tilde{T}^{LB}$ and $\tilde{T}^{UB}$ are defined as follows:

- **Unemployment with permanent income transfers**: The lower bound ($\tilde{T}^{LB}$) closely follows the receipt of UC. It requires an individual to receive UC within 1 month after the end of employment and continue to receive UC with intermediate gaps of less than 4 weeks. If such an intermediate gap or the gap between the end of UC receipt and employment is longer than 1 month, we consider this as an exit to an unknown destination (o) since these exits encompass exits to out of labour force as well as exits to self-employment.

- **Nonemployment**: The upper bound ($\tilde{T}^{UB}$) closely follows a non-employment definition. It requires at least one receipt of UC after an employment spell, but does not impose further restrictions. The resulting spells of unemployment are considered as exits to an unknown destination (o) only if an individual does not exit to employment until the end of the observation period.

By construction of the two unemployment definitions, we observe more UB spells[9] than LB spells because only the lower bound $\tilde{T}^{LB}$ conditions on the receipt of UC within four weeks after the end of employment. In order to avoid a potential sample selection issue if the excluded spells are not random with respect to the treatment effect, we extend the LB sample to match the size of the UB sample as in Lee and Wilke (2005). We do so by adding the missing UB spells to the LB spells. These spells have an observed unemployment duration of zero days and are considered as an exit to an unknown destination state o. This way of treating the added spells is in line with the lower bound definition.

[9]We refer to UB (LB) spells or sample for the sample of unemployment spells that result when applying $\tilde{T}^{UB}$ ($\tilde{T}^{LB}$).

For both unemployment proxies, right-censoring occurs in the case of continued UC receipt at the end of the observation period. For all unemployment spells that exit to employment, the IAB-R01 allows for identifying the location of the new workplace disaggregated to the level of microcensus regions. Thus, by comparing the previous and the new workplace location, it is possible to distinguish local from non-local exits to employment. In the following analysis, a movement between non-adjacent labour market regions (*Arbeitsmarktregionen*) is considered as migration. The 227 labour market regions (LMRs) in Germany comprise typical daily commuting ranges such that for the majority of individuals both residence and workplace are located within the LMR. Since individuals living at the fringe of an LMR may nevertheless easily commute to the adjacent LMR, what is considered a local job change has been extended to include all adjacent LMRs. Finding employment in a non-adjacent LMR should thus necessitate residential mobility in most cases. For each spell of unemployment, the analysis thus distinguishes exits to a local and a non-local job after migration from exits to other destination states. We present estimates for finding local employment and migration only because changes for all other pooled exit states are not easily interpretable and the focus of this application is on transitions to employment.

For our analysis, we include inflow samples for a pre- and a post-reform era. Due to the implementation of stricter sanction rules in 1994, extending the pre-reform era beyond 1995, might mix different reforms. We therefore consider an unemployment spell starting between 1995 and 1996 as a a pre-reform spell. The post-reform era is predetermined by the fact that the implementation of new UNB regulations did not start before 1999. The post-reform inflow sample thus consists of all unemployment spells starting in 1999 or 2000. Since the observation period of the IAB-R01 ends on 12/31/2001, the duration of a post-reform spell is between one and three years only.

Choosing the treatment and control group The aim of the analysis is to identify the effect of being eligible for an extended UNB duration on the transitions from unemployment to either a local job or a non-local job by using the natural experiment that is provided by the reform of the Employment Promotion Act in 1997. In particular, eligibility to an extended UNB duration of more than 12 months was cut for individuals aged 42-44 years, while the PUNBD of individuals below this age was unaffected by the reform. Thus individuals aged 36-41 years serve as the group to control for changing labour market conditions as well

as changing sanction rules when comparing transitions to local and non-local employment before and after the reform. Since only individuals with long UNB entitlements are affected by the reform, the exact choice of treatment and control group has to be conditioned not only on age, but also on the entitlement length at the beginning of the unemployment period. Moreover, the chosen selection rule should be the same for both treatment and control group to ensure that the groups are comparable with regard to their working history. Choosing, for example, all individuals with a maximum duration of UNB receipt in their respective age group results in a non-comparability of individuals in the control and treatment group as the criterium to reach this maximum entitlement is less strict for the younger cohort (see Table 4.2). For this reason, we compute counterfactual UNB entitlements in addition to the actual UNB entitlements at the beginning of each unemployment spell. Both information have to be computed based on the known employment history, age and the known regulations and changes across time (see Appendix A for details). Since the working history for individuals from eastern Germany is not known before 1991 which aggravates the comparability of computed entitlement length, we restrict the analysis to individuals from western Germany. For the counterfactual UNB entitlements, we calculate the entitlement length in the absence of the 1997 reform had the individual been aged 42-44 at the time of benefit claim. As can be seen in Table 4.3, the resulting counterfactual UNB entitlements are quite comparable for both age groups (Pearson chi2(9) = 14.9).

For the subsequent analysis, we choose all unemployment spells that begin with a receipt of unemployment benefits and whose counterfactual UNB duration exceeds 12 months. Moreover, we condition on previous full-time employment to keep the sample more homogeneous in terms of labour force attachment. We also exclude unemployment spells of women because missing information on marital status and dependent children in the IAB-R01 aggravates the interpretation of corresponding results. For the chosen control group and the post-reform treatment group the estimated actual entitlement length as shown in Table 4.4 that is subject to the 1997 reform and the true age of the individual is up to 12 months[10] only. In the pre-reform era, the treatment group is entitled to 18.5 months of UNB receipt on average while in the post-reform era, this average UNB duration falls to 11.8 months.

[10]For some individuals who do not fulfill the criterium for the maximum entitlement length, but still pass the selection criterium, the true UNB duration may be lower than 12 months.

This latter UNB receipt is almost exactly the UNB duration for the control group in the pre- and post-reform era. The average treatment thus is a reduction of UNB entitlements of 6.7 months with the treatment ranging from a reduction of one to a reduction of ten months for individuals aged 44 with maximum UNB entitlements.

Table 4.3: Estimated counterfactual UNB entitlement length for unemployment spells in the pre- and post-reform era by age group[a], IAB-R01

	Age 36-41		Age 42-44	
UNB duration	# spells	%	# spells	%
$\leq$ 2 months	1,226	6.8	551	7.6
3-4 months	1,212	6.7	473	6.6
5-6 months	1,061	5.9	423	5.9
7-8 months	1,092	6.1	440	6.1
9-10 months	1,194	6.6	463	6.4
11-12 months	1,017	5.6	444	6.2
13-14 months	1,182	6.6	486	6.7
15-16 months	1,026	5.7	435	6.0
17-18 months	9,008	50.0	3,505	48.6
Total	18,018	100.0	7,220	100.0

[a] Includes all previously full-time employed individuals born in West Germany whose unemployment spell starts with the receipt of unemployment benefits.

The chosen selection rule for the treatment and control group should ensure some comparability with respect to the working history that builds up claims to UNB. As the working history strongly shapes labour market outcomes, this is quite important in order to minimise selection biases. For a DID approach to be valid, both groups should be comparable in both observed and unobserved characteristics which are likely to affect labour market outcomes. For the available information and some major indicators that can be calculated based on the employment history, Appendix B shows that treatment and control group are quite comparable in most characteristics. Unfortunately, characteristics such as the marital status and dependent children which are likely to affect the likelihood of migration are missing. The subsequent analysis thus rests on the assumption that the composition of treatment and control group in the pre- and post-reform era are as comparable with respect to these unobserved characteristics as they are with respect to the observed characteristics. Another assumption of the DID approach is that both treatment and control group experience similar changes in labour market conditions in the post- compared with the pre-reform era. This

assumption could fail if older workers face more problems to exit unemployment in times of economic downturn than their younger counterparts because of stricter employment protection for older workers. This might be relevant as the post-reform era is characterised by slightly improving labour market conditions, while the pre-reform era rather falls into a period of economic downturn (Bundesanstalt für Arbeit, 2001). On the other hand, stricter employment protection for individuals above 40 only applies occasionally and generally requires a job tenure of more than ten years. For jobseekers between 36 and 44, employment protection should thus be quite comparable and the better labour market conditions in the post-reform era should boost the transition to employment for both groups to a comparable extent.

Table 4.4: Estimated actual UNB entitlement length for unemployment spells with counterfactual UNB entitlements of >12 months in the pre- and post-reform era by treatment and control group, IAB-R01

	Control group		Treatment group	
UNB duration	pre-1997	post-1997	pre-1997	post-1997
6-8 months	2.1%	1.3%	0.00%	1.5%
9-11 months	6.9%	4.5%	0.00%	5.6%
12 months	91.0%	94.2%	0.00%	93.0%
13-14 months	0.0%	0.0%	8.1%	0.0%
15-16 months	0.0%	0.0%	7.2%	0.0%
17-18 months	0.0%	0.0%	56.4%	0.0%
19-20 months	0.0%	0.0%	3.1%	0.0%
21-22 months	0.0%	0.0%	25.2%	0.0%
Average months	11.8	11.9	18.5	11.8
Total spells	4,294	3,577	1,557	1,436

Table 4.5 shows exit types and median unemployment duration for both unemployment definitions. Due to the end of the observation period, the degree of censoring is more pronounced in the post-reform year. Moreover, exits to other destination states are much more likely for the LB spells. The simple descriptive statistics for both unemployment definitions suggest that the treatment group has a somewhat longer unemployment duration, but that the gap between treatment and control group almost disappears after the reform. Among the non-censored spells, both unemployment definitions suggest that the treatment group in the post-reform period almost catches up with the higher exit probability of the control group, especially for exits to local employment.

Table 4.5: Descriptive summary of full sample, IAB-R01

	Control group		Treatment group	
	pre-1997	post-1997	pre-1997	post-1997
LB spells				
median duration (days)	79	73	88	73
exit to local job	54.5% (54.9%)	53.8% (57.0%)	53.4% (53.9%)	52.9% (56.2%)
exit to non-local job	7.7% (7.8%)	8.4% (8.9%)	6.7% (6.8%)	8.2% (8.7%)
exit to other destination	37.1% (37.3%)	32.1% (34.1%)	39.0% (39.3%)	32.9% (35.0%)
total exits	99.3% (100.0%)	94.3% (100.0%)	99.1% (100.0%)	94.0% (100.0%)
UB spells				
median duration (days)	161	124	185	130.5
exit to local job	75.1% (77.4%)	65.9% (74.6%)	72.1% (75.1%)	64.8% (74.1%)
exit to non-local job	12.8% (13.1%)	11.0% (12.5%)	12.5% (13.0%)	11.2% (12.8%)
exit to other destination	9.2% (9.5%)	11.4% (12.9%)	11.4% (11.9%)	11.5% (13.1%)
total exits	97.1% (100.0%)	88.3% (100.0%)	96.0% (100.0%)	87.5% (100.0%)
Total spells	4,294	3,577	1,557	1,436

Cumulative incidence in the pre- and post-reform era Figure 4.4 shows the cumulative incidence curves for exits to local and non-local employment by treatment and control group in the pre- and post-reform years. First of all, note that exit probabilities increase for all groups and exit types in the post-reform years. As has been discussed previously, this may reflect a combination of better labour market conditions compared to the pre-reform years as well as the stricter sanction rules that applied to both the control and the treatment group. More importantly, the figures suggest that the increase in both the local and the non-local exit probability via migration is more pronounced among the treatment group, especially as we reach the end of the shortened UNB duration of 12 months. For exits to local employment, we even observe a small kink for the treatment group after approximately one year of unemployment in the post-reform years. In the pre-reform years where UNB durations have been above 12 months for the treatment group, no such kink can be detected. This timing of events is indicative for a causal relationship of these observed changes to the shortening of UNB receipt in the post-reform years. For exits to non-local employment no clear kinks can be detected, but the increase in the probability of non-local exits seems to be strongest as we reach one year of unemployment.

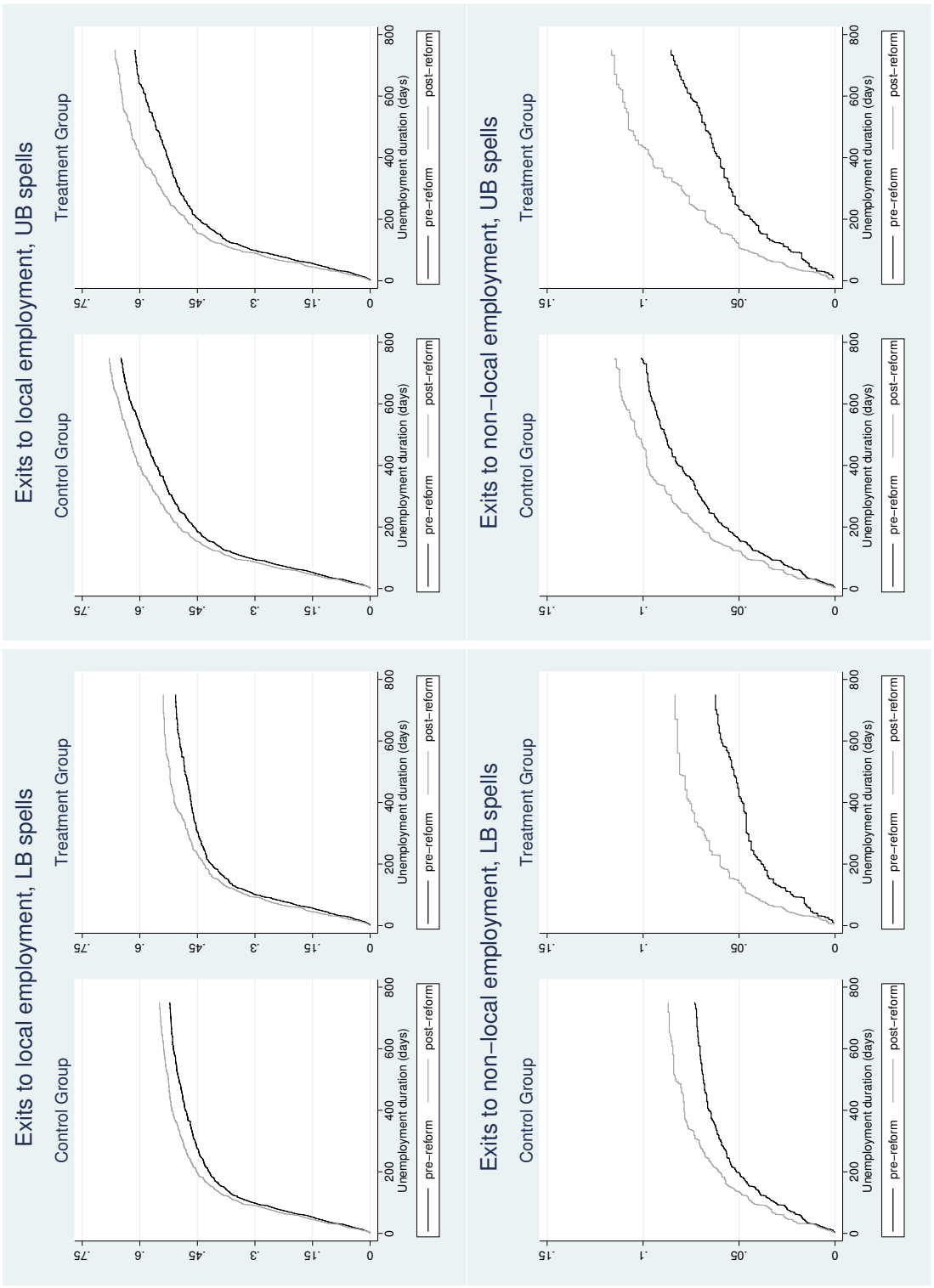
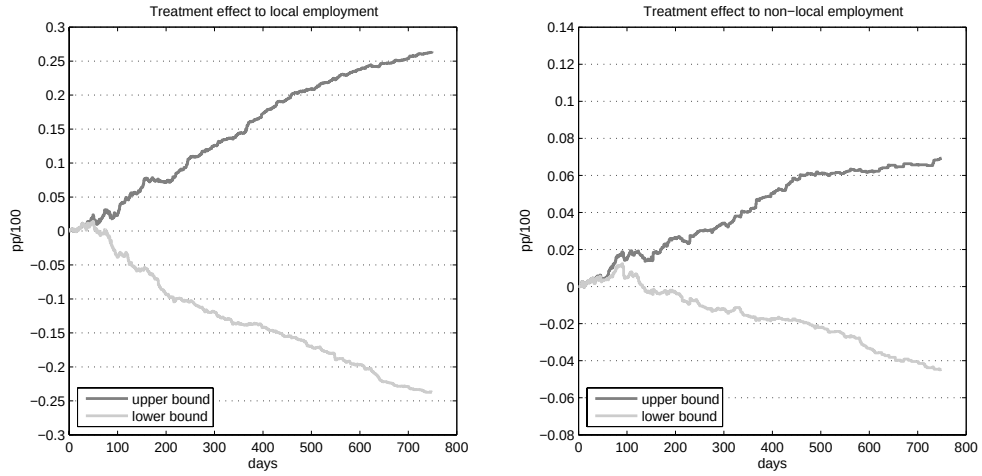


Figure 4.4: Cumulative incidence curves in the pre- and post-reform era by treatment and control group for both unemployment definitions, IAB-R01

In particular, the cumulative incidence curves for non-local exits in the pre-reform era appear flatter after one year of unemployment for the treatment than for the control group. In the post-reform era, the cumulative incidence curve of migration is as steep or even steeper for the treatment compared to the control group. Thus, this may also indicate some positive effect of the shortening of UNB receipt for the treatment group on the observed probability of migration.

Bounds Analysis I As discussed previously, neither of the two unemployment definitions reflects the true unemployment duration. As a consequence, the effects implied by Figure 4.4 do not capture the true effect of the reform on the cumulative incidence of local and non-local employment. We therefore apply formulas (4.17)-(4.18) to bound the true effect of the reform by the cumulative incidence curves for these two unemployment definitions. As a first interesting observation, we find that the resulting bounds do not coincide with the point estimates for the lower and upper bound of the latent variable. As shown in Appendix E, point estimates for the different definitions of the unemployment duration data do not span the full width of our estimated bounds. This suggests that a sensitivity analysis based on different failure time definitions alone may be misleading.

Figure 4.5: Lower and upper bound of treatment effect on the cumulative incidence of local and non-local exits to employment, IAB-R01



As regards the reform effects on local employment and migration, Figure 4.5 suggests that the data insecurity involved due to the unobserved duration until exiting unemployment to one of the unknown exit states precludes any clear inferences. At the very beginning of

unemployment, the upper bound for both exit types is above zero which suggests the same direction of effects as in Figure 4.4. However, after 50-100 days of unemployment bounds are too wide to deduct any treatment effect on exits to local or non-local employment. One reason for the bounds to be so wide is that the natural lower and the natural upper bound for the definition of unemployment that we apply result in many lower bound spells of length zero. These spells are generally uninformative and result in large differences between the distribution of $\tilde{T}^{LB}$ and $\tilde{T}^{UB}$.

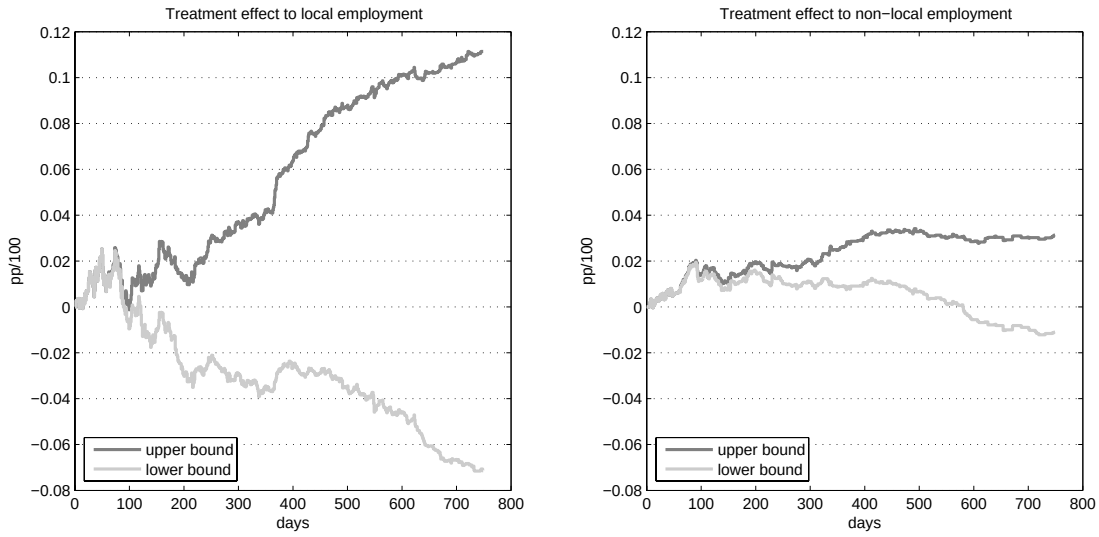
Bounds Analysis II As an approach to tighten the bounds, we impose an additional assumption. Instead of expanding the LB spells to match the size of the UB sample, we restrict the analysis to the spells that are included in both definitions. Given the definitions for $\tilde{T}^{LB}$ and $\tilde{T}^{UB}$ this means that we restrict the analysis to unemployed individuals who receive unemployment compensation within one months after the end of employment. This approach is only valid if the exclusion of spells with a later start of unemployment benefits is a random sample in the sense that the joint distribution of the true unemployment duration remains unchanged. Intuitively, this implies independence of the treatment effect with respect to the exclusion of these uninformative spells.

Table 4.6: Descriptive summary of restricted sample, IAB-R01

	Control group		Treatment group	
	pre-1997	post-1997	pre-1997	post-1997
LB spells				
median duration (days)	107	93	118	95
exit to local job	68.4% (68.9%)	65.2% (70.0%)	66.9% (67.6%)	64.7% (69.8%)
exit to non-local job	9.7% (9.8%)	10.2% (11.0%)	8.5% (8.5%)	10.1% (10.9%)
exit to other destination	21.1% (21.3%)	17.8% (19.1%)	23.6% (23.8%)	18.0% (19.4%)
total exits	99.2% (100%)	93.2% (100.0%)	98.9% (100.0%)	92.7% (100.0%)
UB spells				
median duration (days)	117	99	126	104
exit to local job	78.0% (79.7%)	69.4% (76.4%)	75.5% (78.0%)	69.3% (76.2%)
exit to non-local job	12.1% (12.4%)	11.1% (12.2%)	11.4% (11.8%)	11.1% (12.2%)
exit to other destination	7.7% (7.9%)	10.3% (11.3%)	9.9% (10.2%)	10.6% (11.6%)
total exits	97.9% (100.0%)	90.8% (100.0%)	96.9% (100.0%)	91.0% (100.0%)
Total spells	3,426	2,952	1,243	1,174

As can be seen in Table 4.6, using the restricted sample of spells yields similar descriptive patterns than before. Moreover, the cumulative incidence curves in the pre- and post reform era for the restricted sample in Appendix C indicate comparable shifts than for the full sample. This suggests that restricting the sample to spells with UC receipt within one months does not considerably alter the treatment pattern. A selection issue thus does not seem to be of major concern and the introduction of the sample restriction may be a valid way of tightening the bounds. Bounds for the restricted sample tend to be tighter because the distribution of T^{LB} and T^{UB} are more similar after eliminating one major source of data insecurity by assumption.

Figure 4.6: Lower and upper bound of treatment effect on the cumulative incidence of local and non-local exits to employment, restricted sample, IAB-R01



The resulting bounds in Figure 4.6 are tighter and indicate a positive reform effect for observed transitions to non-local employment. After one year of unemployment, the cumulative incidence of migration for individuals entitled to 12 months of UNB is $1 - 3pp$ higher than for individuals entitled to 16.8 months of UNB on average. In light of the institutional design in Germany, this finding is quite plausible as the counteracting resource effect suggested by Tatsiramos (2003) is likely to be small. This is because unemployed individuals irrespective of whether receiving UNB or UNA get financial support for search costs and moving costs. The negative effect of higher reservation wages in case of higher UNB receipt should thus likely exceed any resource effect. Given the unresolved identification issue of

the competing risks model, however, Figure 4.6 may only be considered as some tentative evidence that unemployment benefits reduce migration. Moreover, despite tighter bounds, the effect of shortening the receipt of UNB on the observed transitions to local employment still cannot be identified from the data and the effect on observed non-local exits remains rather small. One reason for this weak finding may be that due to the institutional design the threat of entitlement loss from a reduction of the PUNBD is likely to be large for a rather small group only. In the final section, we therefore take a look at the heterogeneity of the treatment effect.

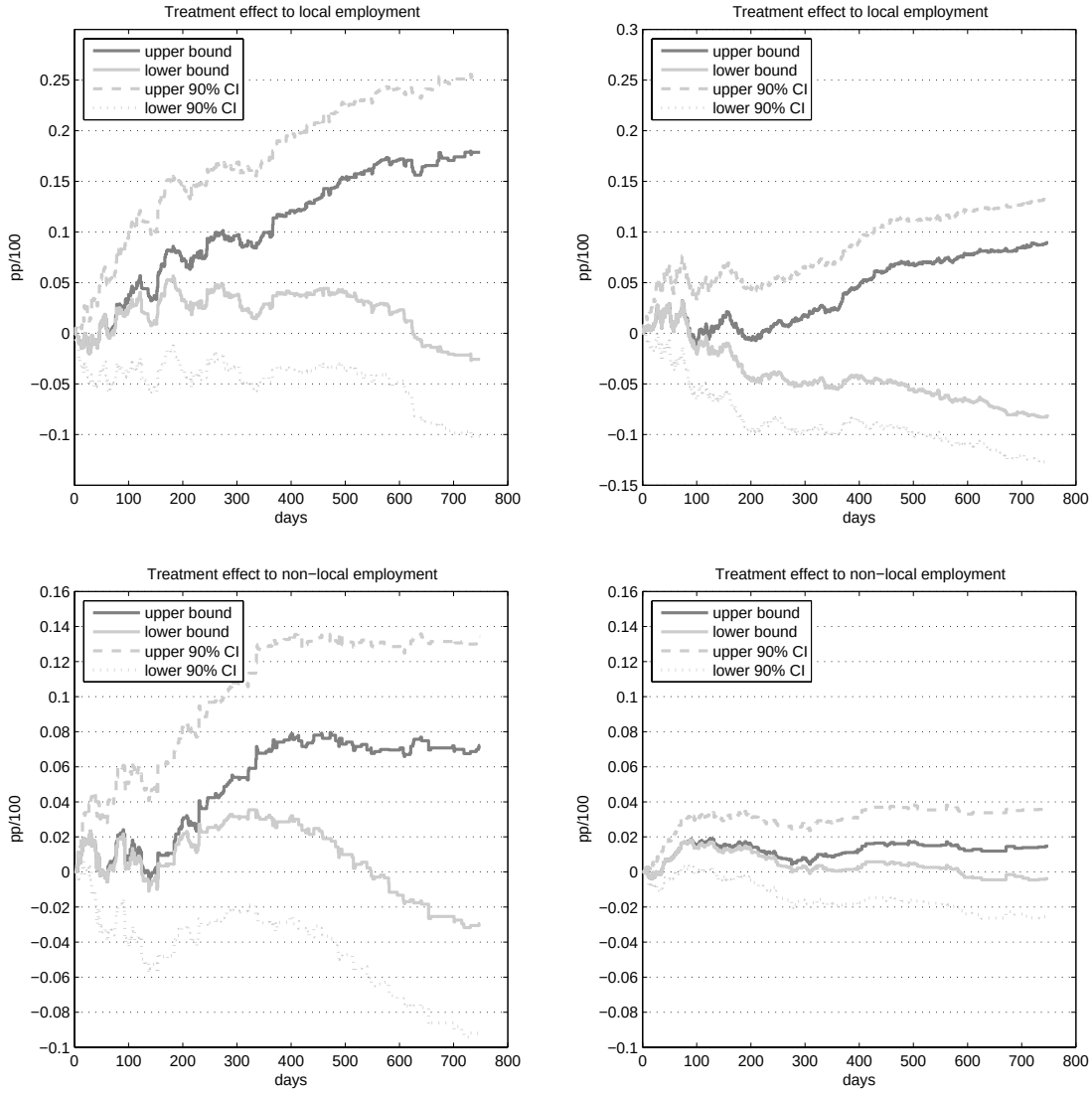
Heterogeneous treatment effects As discussed before, the treatment effect of the reform is unlikely to be homogeneous. Individuals with complementary social benefits are not really affected by the length of UNB receipt. Moreover, individuals who pass the means test for the receipt of UNA only loose around 10% of there former wage income so that the impact of the reform should be limited. Individuals with other financial resources loose the entire unemployment compensation after exhausting unemployment benefits. Unfortunately, the IAB-R01 does not include enough information to actually distinguish between these three groups as the receipt of complementary social benefits is unknown. The wage information included in the IAB-R01 is only a rough indicator of complementary receipt of social benefits as the receipt of social benefits strongly depends on the household context which is unobserved in the IAB-R01. We therefore decided to compare two different skill groups instead because education is highly correlated with wage income and should thus capture some of the aforementioned differences. Less-skilled[11] workers are more likely to receive complementary social benefits or pass the means test for the receipt of UNA than their high-skilled[12] counterparts. The reform effect is thus likely to be weaker for less-skilled individuals. Moreover, distinguishing between skill groups gave a clearer picture compared to looking at different wage quintiles as the latter results were not monotone across wage quintiles. Figure 4.7 therefore presents the findings by skill group. Moreover, we also add the asymptotically valid 90% joint confidence intervals for upper and lower bounds, which are computed following the bootstrap procedure of Horowitz and Manski (2000) and Lee

[11]Includes individuals who are either unskilled or have a vocational training and work as blue-collar workers.

[12]Includes individuals with a tertiary education or white-collar workers with at least a vocational training.

and Wilke (2005). Sample sizes for the two skill groups can be found in Appendix D.

Figure 4.7: Lower and upper bound of treatment effect on the cumulative incidence of local and non-local exits to employment among high-skilled (left) and less-skilled (right) unemployed, restricted sample, IAB-R01



Despite the fact that none of the estimated lower bounds in Figure 4.7 is significantly above zero, the bounds are suggestive for a stronger reform effect on observed exit probabilities for the high-skilled segment for whom the threat of entitlement loss after exhausting UNB is likely to be largest.[13] The point estimates for the bounds indicate that the observed

[13]Due to the identification issue comparisons between groups are not unproblematic as both groups may experience different correlations between exit types which then also affect the resulting cumulative incidence curves.

post-reform probability of migration after one year of unemployment is $3-7pp$ higher for high-skilled individuals while the corresponding bounds for less-skilled individuals suggest a change of $0-1pp$ only. Moreover, point estimates also indicate an increasing observed transition probability to local employment after one year of unemployment of $2-8pp$ for high-skilled individuals only. For high-skilled individuals, the corresponding percentage change on the observed probability of migration is approximately $15-35\%$ while the corresponding change for exits to local employment is around $5-20\%$ only. Under the assumption of independent exit risks, these findings would have a causal interpretation in the sense that extensive unemployment benefits mainly allow for avoiding or postponing migration such that the reduction of UNB entitlements primarily fosters the willingness to migrate. Due to the missing statistical significance which is probably due to the small sample size, however, all these findings are only weakly suggestive for some reform effects on leaving unemployment locally or non-locally and thus call for additional future research with a larger sample size[14].

4.5 Conclusion

This paper has presented an approach that allows for analysing competing failure types in the case of partially missing information concerning the failure times. Partially missing data may occur whenever the state of an individual is partially unobserved such as in the case of unobserved periods in an individual's employment trajectory in administrative individual data. The nonparametric bounds analysis presented in this paper is thus a highly relevant approach for applied researchers who face similar data limitations. It extends the nonparametric bounds analysis for the single risk framework by Abadie (2005) and Lee and Wilke (2005) to a competing risk setting by deriving bounds for the risk specific cumulative incidence curve (CIC). One major advantage of the CIC compared to the marginal survivor curve is that it is still well defined in the case of dependent competing risks. In a simulation, we have demonstrated that this important property of the CIC also carries over to our bounds framework. Although our approach does not resolve the non-identifiability of competing risks and thus precludes a direct causal inference, it provides a flexible descriptive tool for the observed distribution of competing failures. In particular, our approach is fully

[14]The original sample size of the IAB-R01 is in fact quite large with a 2% sample of all individuals working in a job subject to social insurance contributions. The necessary sample selection of treatment and control group, however, reduces the large sample to around 6,000 unemployment spells.

nonparametric in the sense that we do not impose assumptions that may be violated in the real world. In an empirical application of our bounds framework, we have explored the effect of reducing the maximum receipt of unemployment benefits on the observed transitions to either local or non-local employment via migration. For this purpose, we use the variation of unemployment benefit entitlements that is provided by the 1997 labour market reform. Despite avoiding basically any identifying assumption in our bounds framework, we still obtain a number of interesting observations:

- Without showing statistical significance, the bounds are weakly suggestive for a positive effect of reducing the maximum receipt of unemployment benefits on the observed migration probabilities of high-skilled unemployed for whom the threat of entitlement loss after exhausting UNB is likely to be largest. This finding is in line with the previous results from chapters (2) and (3) that entitlements to a long receipt of unemployment benefits tends to prolong unemployment and reduce migration.

- Under the assumption of independent competing risks, the treatment effect on migration clearly seems to exceed the positive treatment effect on exits to local employment in relative terms. This may suggest that extensive unemployment benefits mainly substitute for migration.

- In light of our findings, the current labour market reform in Germany (*Hartz IV*) is likely to foster migration and to accelerate exits to local employment among those for whom the threat of entitlement loss increases. First, the introduction of the means-tested social benefits II (SBII) decouples unemployment compensation after exhausting unemployment benefits from former wage income. This increases the threat of entitlement loss only for those unemployed for whom the former unemployment assistance was more generous than the new SBII. In addition, the reduction of the maximum receipt of unemployment benefits for individuals above the age of 45 since 2003 is also likely to increase transition probabilities to both local and non-local employment. However, the transferability of our results to older age groups may be limited. For older age groups, more restrictive unemployment benefits may rather reduce early retirement, i.e. reduce the inflow into unemployment.

- We generally observe a smooth variation of the bounds with unemployment duration.

This does not suggest any discontinuities in the hazard rates or survivor functions and thus supports the results of the non-stationary job-search theory of van den Berg (1990).

- As another interesting observation, we obtain that point estimates for the lower and upper bound of the latent variable do not span the full width of our estimated bounds. Therefore, a sensitivity analysis based on different definitions of the unemployment duration data alone may be misleading.

The limitations of our approach point towards some interesting future research needs:

- With regard to data limitations, data with more information on individual and household characteristics would be desirable to reexamine our empirical results. Such additional information would also allow to distinguish groups for whom a shorter receipt of unemployment benefits implies different entitlement losses. Moreover, repeating the analysis with a longer post-reform period or a larger sample size should be worthwhile to improve the statistical significance of the data.

- Due to the unresolved identification problem of the competing risk data, the causal inference from our empirical results is limited. Strictly speaking, our results can be interpreted causally only under the assumption of independent risk.

A promising route for future research thus is to combine our bounds framework for partially missing data with attempts to break the non-identifiability of dependent competing risks such as Honoré and Lleras-Muney (2006). However, as a disadvantage to our current bounds framework for cumulative incidence curves, such attempts necessitate additional assumptions.

Appendix

A - Computation of actual and counterfactual UNB entitlements

The entitlement length at the beginning of the unemployment spell is not included in the data and has to be computed based on the known employment history, age and the known regulations and changes across time. For this purpose, we compute the claim period which encompasses a maximum of three years prior to making the UNB claim, but ends with a previous UNB claim within this three years period. In the same token, we calculate the employment duration within the relevant extended claim period of up to seven years prior to making the claim. As previously mentioned, UNB entitlements depend on the duration of socially insured employment within the relevant claim and the relevant extended claim period. Unemployment benefits exceeding six months necessitate at least 12 months of socially ensured employment within the claim period. Thus, an individual with at least 12 months of socially ensured employment within the claim period and 24 months within the extended claim period gets 12 months of UNB. If there is a shortened claim period due to a previous UNB claim, the new UNB claim based on the employment periods after this last unemployment period may be extended up to the age-specific PUNBD by remaining entitlements at the end of the previous unemployment period if the beginning of the last UNB claim lies within the last seven years.

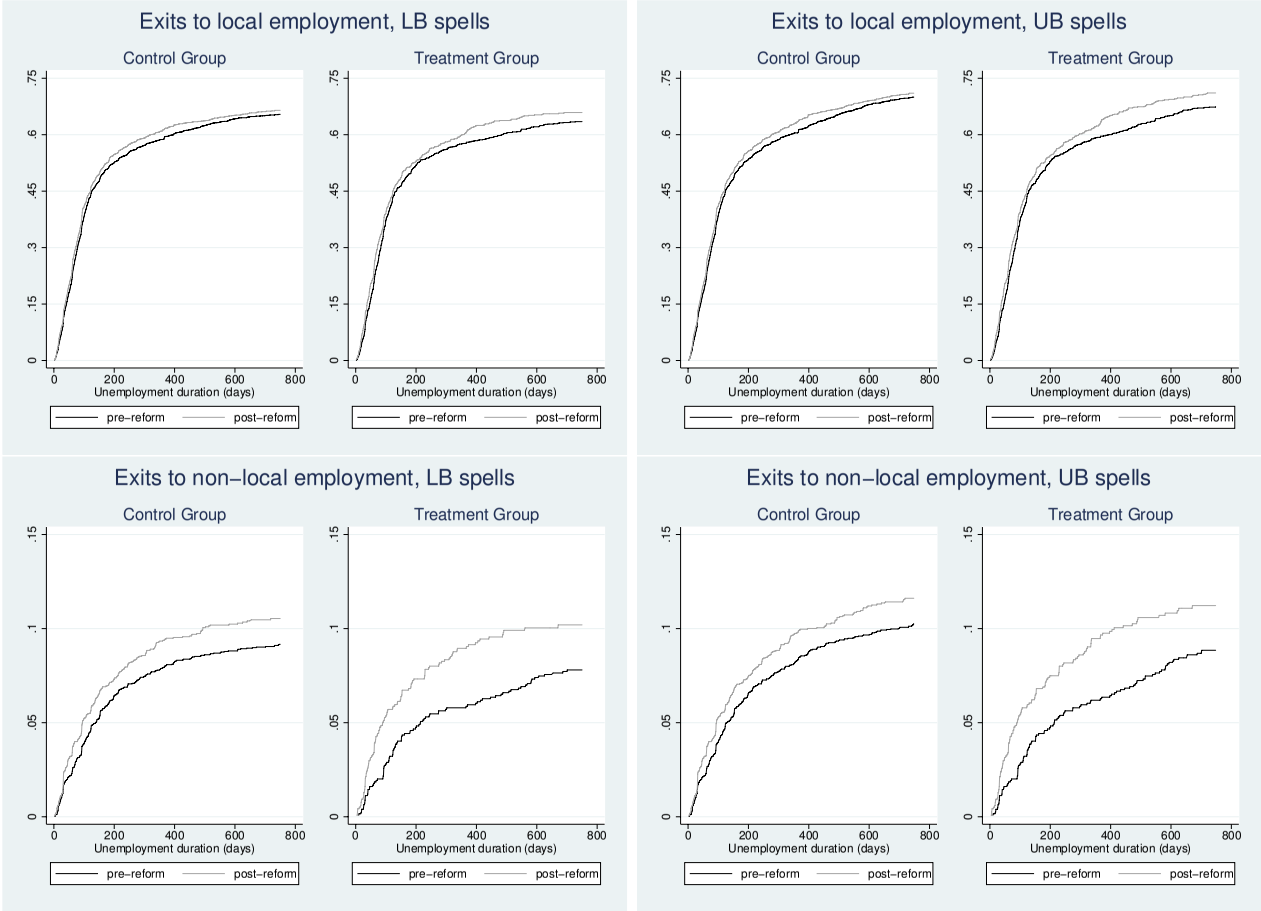
For the estimation of actual UNB entitlements all changing regulations throughout the 1980s and 1990s have been applied. For the counterfactual UNB entitlements, we apply the pre-reform conditions to the post-reform period and compute the UNB entitlements as if all individuals had been 42 by the time of the benefit claim. More precisely, we adjust the whole age history of an individual as if, for example, an individual aged 38 at the beginning of the unemployment period had always been four years older. This adjustment alone does not ensure the comparability of the resulting counterfactual entitlements for the pre- and post-reform period because entitlements depend on the entire work history which is subject to all previous changes in regulations. We therefore compute the counterfactual entitlements for the post-reform period had all changes in regulations been shifted by five years, the difference between the pre- and post-reform period. This procedure ensures a twofold: (i) the comparability of counterfactual UNB entitlements for all age groups irrespective of whether the unemployment period starts prior or after the reform and (ii) the equivalence of

counterfactual and actual UNB entitlements for the treatment group in the pre-reform era. As a consequence, the treatment group in the pre-reform period with counterfactual UNB entitlements of more than 12 months actually has entitlements of more than 12 months while all others who fulfil this criterium actually receive UNB for a maximum of 12 months only, but are comparable to the former group in terms of their employment history.

B - Descriptive summary of sample characteristics

	Control group		Treatment group	
Age (years)	38.3	38.3	43.0	43.0
High school degree	8.6	8.0	8.9	7.7
Vocational training	82.8	83.1	83.1	82.7
Tertiary education	8.6	8.9	8.0	9.5
1st wage quintile	22.0	22.1	23.9	23.9
2nd wage quintile	26.8	30.3	24.0	29.1
3rd wage quintile	20.3	22.9	20.7	20.2
4th wage quintile	17.0	14.8	16.8	14.0
5th wage quintile	14.0	9.9	14.6	12.8
Tenure prev. job (days)	1172.9	1128.6	1385.0	1244.7
Tenure in claim period (days)	1471.4	1434.3	1563.1	1439.2
Prev. recall	17.0	19.6	17.1	19.8
Skilled blue-collar	43.5	43.3	42.8	43.1
Unskilled blue-collar	32.8	31.8	33.7	31.0
White-collar	23.7	25.0	23.4	25.9
Prev. unemployment	73.8	78.9	70.4	77.3
Total spells	4,294	3,577	1,557	1,436

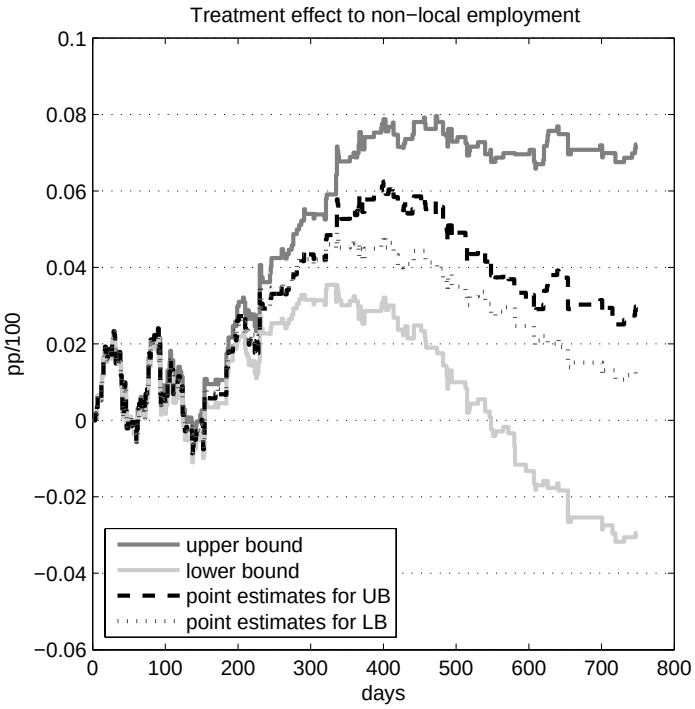
C - Cumulative incidence curves in the pre- and post-reform era by treatment and control group for both unemployment definitions, restricted sample, IAB-R01



D - Descriptive summary of restricted sample by skill group

	Control group		Treatment group	
	pre-1997	post-1997	pre-1997	post-1997
High-skilled individuals				
LB spells				
median duration (days)	154	122	206	122
exit to local job	48.5% (48.9%)	46.5% (49.9%)	43.0% (43.3%)	46.3% (49.2%)
exit to non-local job	19.8% (20.0%)	19.8% (21.2%)	17.2% (17.3%)	18.8% (20.0%)
exit to other destination	30.9% (31.2%)	26.9% (28.9%)	39.2% (39.4%)	29.0% (30.8%)
total exits	99.2% (100.0%)	93.2% (100.0%)	99.4% (100.0%)	94.1% (100.0%)
UB spells				
median duration (days)	200	127	305	154
exit to local job	58.5% (59.8%)	51.8% (56.9%)	53.5% (55.3%)	51.9% (56.0%)
exit to non-local job	25.6% (26.2%)	21.8% (24.0%)	22.3% (23.0%)	21.8% (23.5%)
exit to other destination	13.7% (14.0%)	17.4% (19.1%)	21.0% (21.7%)	19.1% (20.6%)
total exits	97.8% (100.0%)	91.2% (100.0%)	96.8% (100.0%)	92.8% (100.0%)
Total spells	864	815	314	335
Less-skilled individuals				
LB spells				
median duration (days)	99	91	102	92
exit to local job	75.1% (75.7%)	72.3% (77.6%)	74.9% (75.9%)	72.0% (78.1%)
exit to non-local job	6.3% (6.3%)	6.6% (7.0%)	5.5% (5.6%)	6.6% (7.1%)
exit to other destination	17.8% (18.0%)	14.3% (15.3%)	18.3% (18.5%)	13.6% (14.8%)
total exits	99.2% (100.0%)	93.2% (100.0%)	99.7% (100.0%)	92.2% (100.0%)
UB spells				
median duration (days)	105	94	110	95
exit to local job	84.6% (86.4%)	76.2% (83.9%)	83.0% (85.7%)	76.3% (84.5%)
exit to non-local job	7.6% (7.8%)	7.0% (7.7%)	7.8% (8.0%)	6.8% (7.5%)
exit to other destination	5.7% (5.8%)	7.6% (8.4%)	6.1% (6.3%)	7.2% (7.9%)
total exits	97.9% (100.0%)	90.8% (100.0%)	96.9% (100.0%)	90.3% (100.0%)
Total spells	2,562	2,137	929	839

E - Point estimates for lower and upper bound of treatment effect on the cumulative incidence of non-local exits to employment among high-skilled unemployed, restricted sample



The point estimations for the treatment effect using the lower bound and upper bound of the employment duration data are done by using the following formulas:

$$l_{Ir}(t_j|p_{t0},p_{t1},x) = \{I_r^{LB}(t_j|1,p_{t1},x) - I_r^{LB}(t_j|0,p_{t1},x)\} - \{I_r^{LB}(t_j|1,p_{t0},x) - I_r^{LB}(t_j|0,p_{t0},x)\}$$

and

$$u_{Ir}(t_j|p_{t0},p_{t1},x) = \{I_r^{UB}(t_j|1,p_{t1},x) - I_r^{UB}(t_j|0,p_{t1},x)\} - \{I_r^{UB}(t_j|1,p_{t0},x) - I_r^{UB}(t_j|0,p_{t0},x)\}$$

for $r = 1,\ldots,m$ and other notation is the same as in section 4.2.

Chapter 5

What Attracts Human Capital? Understanding the Skill Composition of Interregional Job Matches in Germany

Abstract

By examining the destination choice patterns of heterogeneous labour, this paper tries to explain the skill composition of internal job matching flows in Germany. Estimates from a nested logit model of destination choice suggest that spatial job matching patterns by high-skilled individuals are mainly driven by interregional income differentials, while interregional job matches by less-skilled individuals are mainly affected by interregional differentials in job opportunities. Interregional differentials in non-pecuniary assets slightly contribute to spatial sorting processes in Germany. Such differences in destination choices by skill level are partly modified by different spatial patterns of job-to-job matches and job matches after unemployment. Simulating job matching patterns in a scenario of economic convergence between eastern and western Germany demonstrates that a wage convergence is the most effective means of attracting human capital to eastern Germany.

Keywords: interregional job matches, destination choice, human capital
JEL classification: R23, J61, C35

5.1 Introduction

This paper examines the skill composition of migration flows in Germany by looking at the destination choices of different skill groups. Understanding what attracts skilled migrants to a particular region is of high policy interest because the local availability of a large pool of qualified workers has been considered to facilitate innovative activities and to improve the endogenous growth potential of the region (Lucas, 1988; Romer, 1990). Rauch (1993) and Simon (1998) empirically confirm the positive linkage between the initial human capital endowment of a region and its future economic growth. Similarly, Berry and Glaeser (2005) find that the spatial concentration of human capital in the US increased during the last decades and attribute this to an agglomeration effect that occurs if skilled entrepreneurs tend to create jobs primarily for skilled workers. By contrast, Südekum (2006) stresses skill complementarities as a counteracting force to human capital externalities. For Germany, he confirms a positive link between an initial share of skilled individuals and future regional growth, but attributes this effect to employment growth for low-skilled rather than for skilled workers. Thus, skill complementarities in the production process may result in a more even distribution of human capital across space. However, if the human capital externalities dominate any counteracting skill complementarities, the inward migration of skilled individuals may foster a self-reinforcing spatial concentration of human capital that intensifies regional economic disparities (Nijkamp and Poot, 1997). In the German context, several recent studies suggests that net migration from eastern to western Germany is disproportionately high-skilled (Schwarze, 1996; Hunt, 2000; Burda and Hunt, 2001). This raises strong concerns that a brain drain from eastern to western Germany may reinforce regional east-west disparities[1]. The aim of this paper is therefore to identify major determinants of the skill composition of internal migration flows in Germany. By doing so, the paper provides insights on how policy can promote integration and convergence, a topic that is of high relevance in Germany as well as in a broader European context.

Similar to a recent US study by Hunt and Mueller (2004), the paper considers a number of pecuniary and non-pecuniary forces behind the skill composition of internal migration in

[1] According to Burda and Hunt (2001), the eastern wage level continues to be three-quarters of the western level despite a remarkable wage convergence in the early 1990s. More importantly, the eastern unemployment level is around twice the western.

Germany. This approach fills an important research gap in the European context because existing European studies on the skill composition of internal migration flows in Finland only indicate that high-skilled individuals tend to relocate to high-density urban areas (Ritsilä and Ovaskainen, 2001; Ritsilä and Haapanen, 2003). Wether this is due to a mixture of higher urban wage premia, job opportunities or consumer amenities, however, remains an unresolved question that this paper wants to shed some light on. For this purpose, I use a broad set of regional covariates to capture both pecuniary and non-pecuniary regional disparities. Moreover, based on a sample of job movements between 1995 and 2001, I use a partially degenerate two-level nested logit model to distinguish between job changes within the local area and interregional job changes to one of the destination regions. The paper thus considers migration within a job-changing context by analysing the destination choices of job movers.[2] This enables the spatial pattern of job-to-job matches to be compared with job matches after unemployment which may differ due to different motives for changing a job. Job-to job matches are likely to be mainly voluntary and career-oriented and aim at better job matches. Destinations with good income prospects and attractive amenities may be particularly popular among job-to-job changers. By contrast, job matches after unemployment are more likely to be concerned with job opportunities. To the extent that job-to-job matches and job matches after unemployment are not equally distributed across skill groups, such differences may also affect the skill composition of internal job matching flows. This study therefore extends previous studies by examining differences in destination choices not only by skill level, but also by the type of job match. As another contribution to the literature, the econometric approach of this paper takes account of unobserved interregional heterogeneity. To the extent that amenity valuations differ by skill level, unobserved interregional amenity differentials, for example, may bias the impact of interregional income differentials because observable wage differentials tend to compensate for amenity differentials (Elhorst, 2003). Preceding papers have tended to address this problem by including some amenity indicators such as regional climate differentials (Hunt and Mueller, 2004) which should reduce but not eliminate such biases. Using a pooled sample of job moves between a seven year period

[2] As a drawback, the analysis does not allow for modelling induced job movements. Extending the analysis to endogenously model the probability of changing jobs is not feasible with the data used and is thus left to future research. As a consequence, the study only explains the probability of moving to region k conditional on changing the job (see Bartel, 1979 for a discussion).

allows for including destination and origin fixed effects. This should avoid biases arising from the omission of time-constant region-specific factors (Train, 2002).

Estimation results show some major differences in the spatial pattern of job matches by skill level. Moreover, including destination fixed effects turns out to significantly affect estimation results. In a model with destination fixed effects, the spatial pattern of job matches by high-skilled individuals is mainly driven by interregional income differentials, while job matches by less-skilled individuals are mainly affected by regional differentials in job opportunities. Interregional differences in wage dispersion as well as amenity differentials only weakly contribute to spatial sorting processes in Germany. Moreover, differences in destination choices by skill level are partly modified by different spatial patterns of job-to-job matches and job matches after unemployment. Simulating the spatial pattern of job matches in a scenario of economic convergence between eastern and western Germany thus demonstrates that converging wage levels is the most effective means of attracting human capital to eastern Germany.

The research outline of the paper is as follows. After a short theoretical discussion in section 5.2, section 5.3 and 5.4 introduce the data set and some descriptive evidence regarding the skill composition of internal job matching flows in Germany. Section 5.5 introduces the econometric specification. Section 5.6 discusses estimation results and presents the findings from a simulated economic convergence between western and eastern Germany. Section 5.7 concludes.

5.2 A theoretical underpinning of skill sorting across space

Hunt and Mueller (2004) have recently modelled destination choices as a utility-maximising instead of an income-maximising decision. This approach stresses the role of non-pecuniary returns from moving to a particular region. Similarly, this paper assumes a theoretical framework in which movements between K regions are based on utility maximisation, thus including both pecuniary and non-pecuniary determinants of destination choice. The interesting question now is why such a utility-maximising behaviour induces a sorting of skill groups across space. Despite the rather manageable amount of research on the role of education for destination choices, the literature suggests a number of explanations for such

skill sorting processes. In particular, the net present value of the expected lifetime indirect utility of living and working in k, V_{ik}, should depend on individual i's preferences for certain region-specific attributes as well as his employment and wage prospects in region k. Thus, V_{ik} can be written as

$$V_{ik} = \frac{1}{r}[\alpha_{ik} \int_0^{w^{max}} w \, dF_{ik}(w) + (1 - \alpha_{ik})b_k + a_{ik}] - C_{iok} \tag{5.1}$$

with r as the discount rate, and o denoting the origin region of individual i.

First of all, α_{ik} summarises individual i's chances of finding and keeping a job in region k which may depend on individual i's occupation and skill level and the demand for these characteristics in region k. $1 - \alpha_{ik}$ thus denotes the individual-specific probability of future periods without any wage income but a real transfer income b_k instead that may differ across space due to regional cost-of-living differences. In case of employment, the expected real wage for individual i is given by $\int_0^{w^{max}} w \, dF_{ik}(w)$ which depends on the moments of the wage distribution $F_{ik}(w)$ in region k for individual i's characteristics. While a variance-preserving increase in the mean wage level should attract individuals irrespective of skill level, a change in the wage dispersion may induce skill sorting. According to the extended Roy selection model that applies the Roy model (Roy, 1951) to the international and subsequently to internal migration decisions (Borjas, 1987; Borjas et al., 1992), migrants maximise their income by choosing a destination region that provides the most favourable income distribution for their skill level. In particular, conditional on the mean wage, a high-skilled individual who is likely to draw wage offers from the upper quantile of the wage distribution has a higher expected wage in regions where wage dispersion is greater across skill groups.[3] It follows that high-skilled individuals have incentives to move to regions that reward their human capital investments, whereas less-skilled individuals tend to move to regions with less income inequality in order to reduce the penalty attached to their lack of these skills.

In addition to these pecuniary factors that determine the expected utility of moving to k, a_{ik} captures the value of all non-pecuniary benefits or costs that arise from living in region k. In particular, every location offers a set of natural (e.g. climate), consumer (e.g. the variety of consumption goods and activities) and public goods amenities (e.g. school quality), but also

[3]This only holds if an individual ranks equally in the skill distribution across all regions. In the case of Germany, this assumption may be problematic if formal skills that have been acquired in former East Germany are less valued in western than in eastern Germany.

comes with disamenities (e.g. lack of housing space, pollution, crime rates). Recent research suggests that high-income or educated individuals tend to consume a disproportionate share of consumer amenities and may thus have higher amenity valuations (Brueckner et al., 1999; Glaeser et al., 2001). If amenity valuations rise with skill level, amenity-rich regions should be more frequent destinations for migrants with higher skill levels.

Finally, the costs of moving from the origin region o to region k c_{iok} may be negatively related to human capital (Chiswick, 2000; Brücker and Trübswetter, 2004). This may be a reasonable assumption, if high-skilled individuals are more likely to be compensated for migration costs by their new employer. Migration costs may also be lower due to geographically broader social networks that reduce the information or psychological costs associated with migration. As a consequence, the skill level of internal migration flows might increase with migration distance.

One important insight of this framework is that the proportion of high-skilled individuals moving to k may be affected by skill-specific employment opportunities in region k, the level of amenities, the degree of wage inequality across skill groups and the migration costs involved in moving to region k. In line with these predictions, Hunt and Mueller (2004) find evidence in favour of higher amenity valuations among high-skilled migrants in the US and Canada. Based on a nested logit model of destination choice, their findings also confirm lower migration costs for high-skilled migrants and the implications of the Roy model that high-skilled individuals tend to move to regions with high skill premia.

The objective of this paper is to test these predictions in a German context. As an extension, the paper hypothesises that the skill composition of migration flows may be modified by different destination choices of job-to-job movers and job movers after unemployment since the proportion of job-to-job movers varies across skill groups. For one thing, job-to-job movers may be more likely to make use of career networks and other professional contacts to find a new job. Job-to-job movers may consequently experience favourable job finding conditions α_{ik} even under generally unfavourable job-finding conditions as reflected, for example, in high unemployment rates. By contrast, such general job-finding conditions may be more important for post-unemployment job movers who are less likely to have access to career networks. Secondly, I hypothesise that job-to job matches are likely to be mainly voluntary and career-oriented and aim at better job matches. Destinations with good income prospects

and attractive amenities thus may be particularly popular among job-to-job changers. By contrast, the main migration motive for job movers after unemployment should be to re-enter the labour market whereas regional amenity differentials are likely to be of secondary importance only. The following empirical analysis thus examines destination choices not only by skill level but also by type of job move in order to shed some light on what determines the skill composition of migration flows and thus the allocation of human capital across space.

5.3 Data

The analysis is based on the IAB employment subsample 1975-2001 - regional file (IAB-R01).[4] This register data set contains spell information on a 2 % sample of the population working in jobs that are subject to social insurance payments. As a consequence, the sample does not represent self-employed individuals and tenured civil servants. The data contains spell information on periods for which the individual received unemployment compensation from the Federal Employment Agency (*Bundesagentur für Arbeit*) such as unemployment benefits UB (*Arbeitslosengeld*), unemployment assistance UA (*Arbeitslosenhilfe*) and maintenance payments during further training MP (*Unterhaltsgeld*). Thus, employment histories including periods of transfer receipt can be reconstructed on a daily basis.

Moreover, the IAB-R01 includes the microcensus region of the workplace such that interregional mobility can be identified for job movers by comparing the workplace location of the previous and the current job. Since I only observe workplace locations, any choice of regional boundaries to distinguish between intraregional and interregional mobility entails a possible measurement error if individuals commute across these boundaries. In order to reduce these measurement errors, I define 27 aggregated planning districts (*Raumordnungsregionen*). Planning districts in Germany are defined according to commuting ranges and thus comprise labour market regions that are relatively self-contained. Since using the 97 planning districts for the destination choice model is not feasible, I reduced the number of alternative choices by aggregating planning districts according to an algorithm that reduces the remaining external commuting linkages between these regional planning districts. For details on the procedure see Appendix A. Based on the resulting regional classification, I define the origin and destination region of each job move. According to the definition used

[4]See Hamann et al. (2004) for a detailed description of the IAB-R01.

in the analysis, a job move occurs if there has been a change in employer[5] and the reason for ending the previous spell of employment is denoted as "end of employment". Moreover, no job move is assumed if the next spell of employment indicates the same employer and this new period of employment occurs within 90 days. This restriction ensures that recalls linked to seasonal work are for the most part not counted as job moves.

Distinguishing between job-to-job moves and job moves after unemployment poses some problems. This is because the IAB-R01 does not allow for identifying registered unemployment but only contains information on the receipt of transfer payments. While all unemployed individuals who have previously been employed for at least 12 months are entitled to receive unemployment benefits for a restricted time period, a subsequent time-unlimited entitlement to unemployment assistance is means-tested and thus only applies to individuals who lack other financial resources. This means that it is not possible to distinguish between those who have left the labour force and those who are still unemployed but not receiving unemployment assistance. I therefore distinguish between job-to-job moves and job moves after unemployment by using a proxy for registered unemployment (Fitzenberger and Wilke, 2004; Lee and Wilke, 2005). The resulting types of job moves are defined as follows:

1. **Job-to-job change (JJC):** The job move occurs within 90 days after the last job ended and there has been no intermediate transfer receipt.

2. **Job change after unemployment (UJC):** A UJC occurs if there has been a preceding transfer receipt that terminated less than 90 days before the start of employment. Gaps between previous periods of transfer receipt are no longer than four weeks and transfer receipt started within four weeks after the last spell of employment ended. Since a voluntary job quit entails a suspension of unemployment compensation of at least four weeks, this last restriction ensures that UJC mostly excludes voluntary unemployment.

3. **Job change after all other states (REST):** REST comprises two types of job moves: (1) Job moves without any intermediate transfer receipt but a gap of more

[5]Hunt (2004) suggests that high-skilled individuals are quite likely to be interregionally mobile while staying with the same employer. I deliberately exclude this type of migration because these movements are largely determined by site locations of the employer and not by a decision-making process that considers all alternative locations.

than 90 days between both spells of employment. (2) Job moves with intermediate transfer receipts that does not fulfill the UJC definition due to longer gaps before, during or after transfer receipt. In both cases, long gaps in the employment history may be due to other unobserved labour market states (e.g. self-employment, out-of labour force).

For the subsequent analysis, I use only JJC and UJC since the remaining job moves (REST) are a very heterogeneous sample for whom the intermediate labour market status and thus also the intermediate whereabouts are rather unclear. Moreover, due to the data limitations, the analysis mainly excludes graduates who enter the labour market for the first time unless these individuals have been recorded in the IAB-R01 due to some previous employment relationship that has been subject to social contribution payments. Furthermore, I restrict the sample to job moves occurring between 1995 and 2001 since prior to 1995 there have been dramatic changes in the demarcation of eastern regions that complicate any regional analysis. In addition, I restrict the sample to prime-age males aged 25 to 45 years in full-time employment in order to receive a relatively homogeneous sample. Despite a growing literature regarding the substantial east-west migration of women in Germany (Kröhnert et al., 2006), I exclude women from the analysis due to data restrictions. In particular, the IAB-R01 does not include information on marital status and single and married women cannot therefore be separated. Since these two groups are likely to behave quite differently, with married women often being tied movers, I decided to restrict the analysis to male job movers. All these data-driven sample selections should be borne in mind for the later analysis as they limit the generalisability of the empirical findings to other excluded labour market segments. Still, examining destination choices among male direct job movers and job movers after unemployment should yield some interesting insights into the factors that drive the skill composition of migration flows in Germany. For this analysis, I distinguish between high-skilled job movers with a college or university degree and less-skilled individuals who are either unskilled or have a vocational training.[6] In Germany, unskilled individuals with only a high-school degree comprise less than 10% of all individuals. Based on these defini-

[6]I address the problem of inconsistencies in the education variable in the IAB-R01 by using the IPI imputation rule that has been proposed by Fitzenberger, Osikominu and Völter (2005). This imputation rule assumes that educational degrees do not get lost and that missing values may be overwritten by previous information on the education level if available.

tions, I observe 116,978 JJC and 85,066 UJC by 26,457 high-skilled and 175,587 less-skilled individuals in the period from 1995 to 2001. Moreover, 72% of all individuals experience more than one job move within the seven year period.

5.4 Background and descriptive evidence

In order to give some descriptive evidence regarding differences in spatial job matching patterns by skill level and type of job change, I consider job matching between four macro regions (north, mid, south, east) as shown in Figure 5.1.

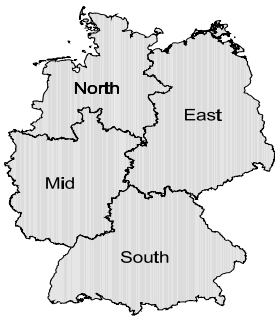


Figure 5.1: Four German macro regions

Table 5.1 shows average economic conditions in these regions between 1995 and 2001. There are strong disparities among the three western regions (north, mid and south) with regard to unemployment rates. While the south has unemployment rates which are much lower than the national average, the north and to a lesser extent the mid are characterised by much higher rates of unemployment. Eastern Germany still lags behind economically with unemployment rates around twice the average rate of the three western regions. Moreover, eastern wages continue to be one-quarter below the western wage level despite a remarkable wage convergence during the 1990s. The observed downward trend in east-west migration from an initial peak in the early 1990s has mainly been attributed to this wage convergence (Hunt, 2000; Burda and Hunt, 2001). Wage dispersion continues to be less pronounced in the eastern than in the western regions despite growing wage inequality in eastern Germany since the 1990s. According to the Roy selection model, this should contribute to a positive selection of east-west migrants.

Table 5.1: Average economic conditions in four German macro regions, 1995-2001

Indicator[a]	**East**	**North**	**Mid**	**South**
Median daily wage in euros	60.2	81.4	83.0	83.7
Wage variance index[b]	0.84	1.17	1.06	1.01
Unemployment rate	17.9	11.0	10.4	7.5
Employment growth in %	-1.5	1.2	1.1	1.7

[a] For details on the data sources and definitions of indicators see Appendix B and C.
[b] An index of < 1 indicates below average wage variance. See Appendix C for details.

Bearing these regional disparities in mind, Table 5.2 shows job matching patterns by origin and skill level between these macro regions. Note that an interregional job move can occur within the same macro region since each of these regions consists of several sub-regions. Consistent with the migration literature, high-skilled individuals are much more likely to experience an interregional move than less-skilled individuals. More importantly, destination choice patterns also differ by skill level. While high-skilled job changers are, for example, two to four times as likely to move to the south than their less-skilled counterparts, the likelihood of moving to the east is similar across both skill groups.

Table 5.2: Mobility pattern by origin and skill level, IAB-R01, 1995-2001

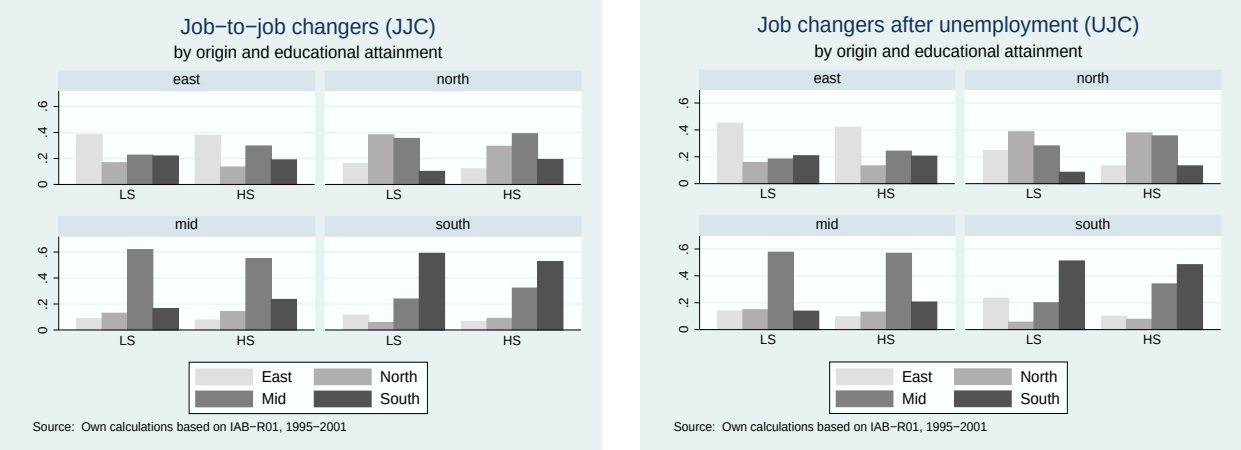
			Destination (in %)				
Origin	**Skill level**	**Obs.**	**Stay Home**	**East**	**North**	**Mid**	**South**
East	Less-skilled	49,935	84.0	6.7	2.6	3.3	3.4
	High-skilled	4,862	69.1	12.1	4.2	8.7	6.0
North	Less-skilled	28,009	82.1	3.5	6.9	5.9	1.7
	High-skilled	3,913	58.2	5.2	13.2	16.0	7.4
Mid	Less-skilled	56,085	79.5	2.1	2.8	12.4	3.2
	High-skilled	10,364	58.4	3.3	5.7	23.0	9.6
South	Less-skilled	41,558	83.3	2.5	1.0	3.8	9.4
	High-skilled	7,318	62.1	2.6	3.3	12.4	19.6

Note: Less-skilled individuals with high-school degree or vocational training; High-skilled individuals with tertiary education.

According to the theoretical framework of the previous section, different destination choices by skill level may partially reflect different spatial job matching patterns of job-to-

job moves and job moves after unemployment since skill groups are not evenly distributed across these types of job mobility. Figure 5.2 thus displays destination choice patterns of interregional job moves not only by skill level but also by type of job move. According to Hotelling test statistics, differences across skill groups remain highly significant after controlling for the type of job mobility. Moreover, destination choice patterns also differ significantly between JJC and UJC when controlling for skill level. This suggests that the skill composition of job matching flows in Germany may also be affected by different job matching patterns of job-to-job changers and job changers after unemployment.

Figure 5.2: Destination choice pattern by skill level, origin and job status, IAB-R01, 1995-2001



Note: LS: Less-skilled; HS: High-skilled

Table 5.3 looks at the resulting skill composition of job matching flows between the four macro regions. In particular, it shows the share of high-skilled individuals among job movers between the four regions. On average, 24.5% of all interregional job moves accrue to high-skilled individuals, but there are large differences in the skill composition of particular migration paths. The skill level of flows to the east and the north, for example, is lower than average, while the skill level of flows to the south and the mid region is above the average. Interestingly, regions with high-skilled inward migration also tend to have high-skilled outward migration and vice versa.

The skill composition of inward and outward flows does not say much about the implied net flow of less-skilled and high-skilled individuals. Table 5.4 therefore looks at net migration flows and the induced net employment change by skill level for the four macro regions.

Table 5.3: Share of high-skilled individuals among interregional job movers between the four regions, IAB-R01, 1995-2001

	Destination				
Origin	**East**	**North**	**Mid**	**South**	**All**
East		13.6%	20.4%	14.7%	16.5%
North	17.2%		27.5%	37.8%	26.5%
Mid	22.5%	27.3%		35.7%	29.8%
South	15.5%	36.8%	26.5%		30.6%
All	18.7%	24.0%	28.6%	28.4%	24.5%

Apparently, both the east and the north experience net losses of human capital. In line with Büchel et al. (2002), the descriptive evidence thus points towards a continued brain drain from eastern to western Germany. Moreover, the east not only loses high-skilled migrants to the south and the mid, but also experiences an even larger net loss of less-skilled migrants. By contrast, the mid and especially the south have positive net flows for both skill groups. For the south, the employment change that is induced by these net flows is larger for high-skilled than for less-skilled individuals.

Table 5.4: Net migration flows and induced net employment change by skill level, IAB-R01, 1995-2001

Region	**Net migration**		**Net emp. change**	
	LS	**HS**	**LS**	**HS**
East	-1447	-183	-1.40%	-1.17%
North	175	-83	0.24%	-1.25%
Mid	337	29	0.20%	0.17%
South	935	237	0.67%	1.75%

Note: Employees by skill level are computed based on the IAB-R01 at the beginning of the observation period (01/01/1995).

The job-matching flows thus indicate a re-allocation of population from the east to the west and a re-allocation of human capital from the east and the north to the south mainly. The descriptive evidence suggests that destination choice patterns differ by skill level and type of job move. The following econometric analysis thus examines destination choices of heterogeneous labour in order to identify the factors behind these observed sorting processes.

5.5 A partially degenerate nested logit model

Following the well-known random utility approach to discrete choice problems (McFadden, 1981), the probability that individual i with origin o chooses destination d can be written as:

$$P_{iod} = P[V_{iod} + \epsilon_{iod} > V_{iok} + \epsilon_{iok}] \quad \forall\, k \neq d \tag{5.2}$$

with V_{ioj} denoting the observed utility for individual i of moving to region j=d,k. ϵ_{ioj} is the unknown stochastic part. Assuming independent, identically extreme value distributed error terms between all destination choices yields the logit specification which has been used by a number of recent destination choice studies (Davies et al., 2001; Schündeln, 2002). Since the simple logit representation is inappropriate if choices are related due to unobserved utility components, I choose a nested logit specification that slightly relaxes the independence assumption of the logit specification by allowing for some correlation among non-origin regions.[7] More specifically, I use a partially degenerate nested logit model that distinguishes between two upper-level branches: staying in the local area (s) and migrating (m). At the lower-level, the branch m distinguishes between all destination regions while for the degenerate branch s, the origin region is the only choice. This model thus allows for the case that all choices that involve residential mobility are related due to some unobserved migration cost, but still assumes independence between all non-origin regions in branch m conditional on all observed factors, i.e. the the Independence of Irrelevant Alternatives (IIA) assumption has to hold with branch m.

The nested logit model can be decomposed into the product of the marginal probability of choosing branch m or s (P_{il} with $l = m,s$) and the conditional probability of choosing alternative k conditional on choosing the branch ($P_{ik|l}$). The conditional probability for the non-degenerate branch m can be written as

$$P_{ik|m} = \frac{exp(\gamma' z_{ik})}{\sum_{k \in m} exp(\gamma' z_{ik})} \tag{5.3}$$

while $P_{io|s} = 1$ for the degenerate branch. γ denotes a parameter vector. z_{ik} are covariates that vary across non-origin regions. The upper level marginal probability of migrating can

[7] A less restrictive multinomial probit that allows for correlations between all alternative choices is infeasible due to the computational burden that results from 27 alternative choices and the large sample size.

be written as follows:

$$P_{im} = \frac{exp(\beta'_m w_i + \zeta_m iv_{im})}{1 + exp(\beta'_m w_i + \zeta_m iv_{im})}. \tag{5.4}$$

with

$$iv_{im} = ln[\sum_{k \in m} exp(\gamma' z_{ik})]. \tag{5.5}$$

β_m is a parameter vector that measures the effect of each individual-level characteristic w_i on the probability of migration. iv_{im} refers to the inclusive value which links the upper with the lower model. In particular, $\zeta_m iv_{im}$ may be interpreted as the expected utility individual i derives from choosing among all non-origin regions, i.e. from migrating. Moreover, the inclusive value parameter ζ_m reflects the degree of independence among all non-origin regions. Since $\zeta_m = 1$ has been rejected for all estimations in the following section, the alternative choices cannot be considered fully independent such that the nested logit model turns out to be an appropriate specification. I estimate a non-normalised nested logit (NNNL) for which the utility of the lower level model has not been rescaled by the inverse of the inclusive value parameter (Daly, 1987). The normalised utility maximising nested logit (McFadden, 1978) is typically preferred for its consistency with utility maximisation if $0 < \zeta_m < 1$. The NNNL specification is consistent with utility maximising behaviour only if no coefficients are common across branches and ζ_m lies within the interval $[0; 1]$ (Koppelman and Wen, 1998; Heiss, 2002; Hensher and Greene, 2002). Since both conditions are fulfilled in the subsequent estimations, using the NNNL specification is a feasible approach. I estimate the NNNL sequentially by estimating the lower level model and the inclusive value before estimating the upper level model. This sequential estimation is less efficient than simultaneous estimation by full information maximum likelihood (FIML). Moreover, due to the inclusive value estimate, the standard errors of the upper level model may be biased downward (Amemiya, 1978). Thus, FIML is clearly preferable but comes at the cost of difficult numerical maximisation since the log-likelihood function is not globally concave. Moreover, FIML was computationally infeasible for the complete sample. Since the main focus of the paper is on lower level estimates for which both point estimates and standard errors are consistent, I therefore decided to use the sequential estimation method. Both point estimates and standard errors for upper level covariates were quite similar when comparing sequential estimates with FIML estimates for some sub-samples. This suggests that the sequential

estimation bias may be negligible. For all estimations, I further impose standard errors that are robust to clustering at the regional level in order to avoid downward biased standard errors (Moulton, 1990).

Upper level covariates w_i consist of individual-level characteristics that may be expected to affect individual mobility decisions. In particular, these covariates encompass age, previous job status, previous sector of activity, previous type of occupation and wage income in the job prior to the job move. Unfortunately, the IAB-R01 does not include important household characteristics such as home ownership and marital status which repeatedly have been shown to affect the propensity to be mobile. Instead, the data set allows for capturing the individual employment history (e.g. previous average tenure, previous recall by the former employer, total duration of all previous non-employment periods) which should at least reduce some of the unobserved heterogeneity among individuals. I also include an indicator for individuals of East German origin to capture differences in the propensities to be mobile that may be related to the different relevance of geographical mobility in both parts of Germany. Moreover, since all observed job moves between 1995 and 2001 are pooled for the analysis and individuals may experience more than one job move during this period, an additional indicator for multiple job moves should control for major differences between multiple job movers and individuals with only one job move. In addition, I include origin fixed effects in order to capture differences in the propensity to be mobile across origin regions as has been shown in Table 5.2. Appendix E contains summary statistics for all upper level covariates.

Lower level covariates z_{ik} vary across non-origin regions and are intended to capture observed utility differences between alternative destinations as suggested by the theoretical framework in section 5.2. As an indicator of regional job-finding conditions for individual i, I use the regional unemployment rate[8], regional employment growth in individual i's skill group and the share of high-skilled employed in region k . While the unemployment rate indicates general job-finding conditions, higher employment growth in individual i's skill group indicates improving employment prospects. Moreover, a region with a high level of qualified jobs as reflected by a high share of high-skilled employees should offer

[8]Unfortunately, no regionally disaggregated unemployment rates by skill group are available.

favourable job-finding conditions for high-skilled job movers. z_{ik} also includes the median wage in individual i's sector of activity as an indicator of interregional differences in the wage level[9]. Moreover, I use the ratio between the 80th and 20th wage percentile in region k as an indicator of the regional wage dispersion across skill groups.[10] According to the theoretical framework, higher wage levels should attract migrants irrespective of skill level while a higher degree of wage inequality should attract mainly high-skilled individuals. In addition to income differentials and job-finding conditions, I also try to capture a number of non-pecuniary regional differences. Unfortunately, there are only few indicators that may be considered to capture such non-pecuniary aspects for which a time series is available for the study period. In particular, I include population levels as a proxy for urban-scale related consumer amenities as suggested by Herzog und Schlottmann (1993). Moreover, I include the population density as a measure of agglomeration effects as suggested by Ciccone and Hall (1996)[11]. While urban-scale related amenities should be attractive for migrants, especially high-skilled ones, a denser agglomeration for a given urban scale may also capture disamenities such as pollution or lack of housing space.[12] In addition, I use hotel capacities per resident as a proxy for the general attractiveness of the region as proposed by Glaeser et al. (2001). Moreover, I include regional child care facilities as an indicator of the availability of a specific type of public goods. In order to capture also a specific source of disutility, I include regional crime rates. Regional land price differentials are used as a proxy for interregional cost of living differentials. In addition, the model includes the distance between origin and destination region as a measure of migration costs that may be related to distance such as psychological costs.

Estimation results based on this specification may be biased if covariates such as employment growth and population size are endogenous due to a simultaneity issue. In order to mitigate this problem, I use lagged values for all covariates z_{ik} for which such a simultaneity

[9]When using the regional wage level across all sectors, estimates turned out to be weaker. Apparently, interregional differences in the sector wage level appear to be more relevant for mobility decisions.

[10]Both income indicators control for different regional compositions of the labour force such that differences in these indicators reflect differences in labour prices only. Appendix C includes a short description of the methodology which is based on Hunt and Mueller (2002).

[11]In fact, Ciccone and Hall use employment density as a measure of agglomeration economies, but population densities should also be an appropriate indicator.

[12]Positive agglomeration effects such as higher productivity levels due to closer proximity of workers and lower transportation cost, should mainly be captured by the regional wage distribution.

issue is likely to arise (see Appendix B). Even lagged values, however, can be endogenous due to the persistence of unobserved regional characteristics over time. For this reason, I include fixed effects for each destination region at the lower level of the model in order to avoid biases from omitting relevant destination-specific factors. Unobserved characteristics of a particular migration path such as the cultural proximity between origin and destination region may, however, continue to bias estimation results. Since it is not possible to include fixed effects for each origin-destination pair, I only include fixed effects for movements across the former inter-German border and for movements between northern and southern Germany. Moreover, in order to measure a possible reluctance of individuals to cross the former inter-German border, the dummy variable for movements from western to eastern Germany only applies to individuals born in West Germany, while the dummy for the opposite direction only applies to individuals born in East Germany. For migration between the north and the south of Germany, the data set does not include the information to restrict the dummies to individuals born in these parts. The dummies thus apply to all individuals living in these regions. Including lagged covariates, regional fixed effects for destination regions and fixed effects for some major migration path should clearly reduce potential biases compared to earlier studies that do not consider any fixed effects such as Hunt and Mueller (2004).

Except for the distance measure, all continuous lower level covariates z_{ik} are defined as differences between the standardised values for the destination and the origin region, i.e. $z_{ik} = \tilde{z_{ik}} - \tilde{z_{io}}$. This reflects the notion that destination choices are typically made by comparing potential destinations with the current region of residence. As a drawback, however, this imposes the restriction that responses to changes in the origin or the destination region are symmetric.[13] Appendix B and C lists the exact definitions and data sources of all lower-level variables, while Appendix D gives the corresponding summary statistics.

Marginal effects Due to defining lower level covariates as differences between standardised values, marginal effects measure the effect of an increase in the difference between origin and destination region by one standard deviation. Thus, marginal effects of a change in z_{ik} on the conditional probability of moving to region k are comparable for these covariates and

[13] A less restrictive specification with origin-specific characteristics in the upper-level model proved quite unstable such that I decided to stick to the more restrictive use of destination-origin differences.

have been computed as follows:

$$\frac{\partial P_{ik|m}}{\partial z_{ik}} = \gamma_z P_{ik|m}(1 - P_{ik|m}) \tag{5.6}$$

For dummy variables, marginal effects have been calculated instead as $\triangle P_{ik|m}/\triangle z_{ik} = P_{ik|m,z_{ik}=1} - P_{ik|m,z_{ik}=0}$. For the upper level model, marginal effects of a change in w_i on the marginal probability of moving to region d are given as

$$\frac{\partial P_{im}}{\partial w_i} = \beta_w P_{im}(1 - P_{im}) \tag{5.7}$$

for continuous covariates and as $\triangle P_{im}/\triangle w_i = P_{im|w_i=1} - P_{im|w_i=0}$ for dummy variables. For both lower and upper level marginal effects, the delta method has been applied to calculate standard errors. Marginal effects and standard errors shown in the subsequent tables always refer to the average effects in the sample population (Train, 2002).

5.6 Estimation Results

Following the sequential estimation procedure, this section discusses the lower level model of destination choice of interregional job moves before briefly discussing the upper level estimates for the decision as to whether to change a job intra- or interregionally. Based on these results, I then examine the implied change in the mobility pattern of job moves in case of an economic convergence between eastern and western Germany.

Lower level estimates Table 5.5 shows estimated marginal effects on the conditional probability of moving to destination k by skill level for the pooled sample of job-to-job moves and job moves after unemployment. Specification A includes neither destination-specific fixed effects nor dummy variables for specific migration paths while specification B includes these additional covariates. Comparing both specifications in Table 5.5 suggests that including a number of regional amenity indicators in specification A does not suffice to prevent biases from unobserved time-invariant interregional amenity variations. In particular, the effect of the wage level seems to be downward biased while the impact of the unemployment rate is upward biased. These biases are consistent with the notion that higher local amenities compensate for lower wages and higher unemployment rates. In this case, the interregional wage differential is negatively related to the unobserved interregional amenity differential, while the unemployment differential should be positively related to the amenity

differential (Elhorst, 2003). Moreover, the amenity indicators also seem to be biased due to unobserved region-specific factors. As an example, specification A indicates that higher crime rates attract individuals, while no such evidence can be found for specification B. To sum up, the findings indicate that estimates without destination fixed effects may be biased. Specification B also seems to be more reliable than specification A when it comes to testing the independence of irrelevant alternatives assumption by running a Small-Hsiao test[14] (Small and Hsiao, 1985) for excluding each of the 27 regions, respectively. Table 5.5 shows how many of these 27 test statistics suggest that the independence assumption is incorrect. The Small-Hsiao test confirms the iia assumption at least for model B for almost all regions. This can be seen as some positive evidence that the nested logit model is well specified. For the subsequent analysis, unless stated otherwise, I restrict the discussion of covariate effects to the more reliable specification B. In order to examine whether the type of job move matters for the spatial pattern of job matches, Table 5.6 thus displays estimation results by skill level and type of job move for specification B only.[15]

Economic conditions As expected, interregional job changers tend to move to regions with higher wage levels in their sector of activity. Interestingly though, the last column in Table 5.5 suggests that this effect is significantly stronger at a 5% significance level for high-skilled than for less-skilled interregional job movers. While for less-skilled individuals a one standard deviation increase in the sector wage level in region k increases the probability of moving to k by only $0.5pp$, the corresponding effect for their high-skilled counterparts is four times as large. Consistent with higher labour supply elasticities among high-skilled as compared to less-skilled individuals[16], high-skilled individuals thus have stronger preferences for high-wage regions. Consistent with the theoretical notion discussed in section 5.2 that income prospects may be more important for career-oriented job-to-job moves than for job

[14] As discussed in Small and Hsiao (1985), the Small Hsiao test should be preferred to other test statistics such as the Hausman test because it avoids the computational and inference problems of the Hausman test that arise if the inversion of the difference between two similar parameter matrices is not positive definite or close to being singular. Indeed, the results from the Hausman test do not seem to be very reliable as the test statistic is often not positive definite depending on sample choice and included covariates. I therefore rely on the Small-Hsiao test instead.

[15] Results for specification A are available from the author upon request.

[16] Arntz et al. (2006a) estimate labour supply elasticities by skill groups for Germany based on the ZEW microsimulation model and find that labour supply elasticities for high-skilled individuals exceed labour supply elasticities for less-skilled individuals.

moves after unemployment, point estimates in Table 5.6 further suggest that the wage level is a more important determinant of destination choice for job-to-job movers than for job movers after unemployment. Differences between the two types of job movers are not significant though (p-value for high-skilled: 0.23).

There is no significant evidence in Table 5.5 that high-skilled job movers prefer regions with a high wage dispersion, while there is some evidence that their less-skilled counterparts avoid such regions. Controlling for the type of job move in Table 5.6 does not alter this result. Consistent with the extended Roy model, this finding may thus be suggestive of some weak skill sorting based on interregional differences in wage inequality. Compared to the U.S. study by Hunt and Mueller (2004), however, the impact of wage inequality is very weak.[17] This may be because interregional differences in wage dispersion are much smaller in Germany than in the US with the exception of east-west disparities. Such disparities, however, may be of minor importance compared to the strong east-west differences in wage levels. In this case, a selection based on interregional differences in wage dispersion may not be a major determinant of the skill composition of interregional job moves in Germany. Instead, interregional wage level differences not only affect the level of inter-state migration in Germany as suggested by Burda and Hunt (2001) but also strongly affect the skill composition of these flows.

The skill composition of interregional job flows is also affected by interregional differences in employment opportunities. More specifically, Table 5.6 shows that irrespective of the type of job move there is significant evidence that less-skilled individuals tend to move to regions with low unemployment rates, while no significant evidence can be found for for their high-skilled counterparts. By contrast, significantly positive effects of employment growth can be found for job-to-job movers only. Consistent with the hypotheses in section 5.2, generally favourable job-finding conditions as reflected by low unemployment levels, are important for less-skilled job movers who are less likely than their high-skilled counterparts to make use of interregional career networks and may thus experience strong job competition in regions with high unemployment levels[18].

[17]The stronger U.S. findings may also reflect specification issues since Hunt and Mueller (2004) do not use standard errors that are robust to clustering at the regional level.

[18]Unfortunately, the unemployment rate by skill group which would be more informative on this issue is not available in Germany on a regionally disaggregated level.

Table 5.5: Lower level marginal effects[a] $\frac{\partial P_{id|m}}{\partial z_{id}}$ by skill level for a pooled sample of JJC and UJC in percentage points (pp), IAB-R01 1995-2001

	Model A		Model B		
Variable	**LS**	**HS**[b]	**LS**	**HS**[b]	p-value[c]
Median sector wage	0.057	1.364**	0.532**	2.061**	0.031
Wage variation	-0.222*	-0.029	-0.382†	0.029	0.239
Unemployment rate	-0.453**	-0.153	-1.265**	-0.591	0.290
Employment growth	0.638**	0.179†	0.130	0.217†	0.572
Share of HS employment	0.703**	1.293**	0.634	-0.573	0.176
Log(Distance)	-6.446**	-4.664**	-6.131**	-4.317**	0.089
Population size	1.271**	1.365**	0.393*	0.483*	0.543
Population density	-0.222**	-0.314*	-0.228*	-0.183	0.589
Crime Rate	0.292**	0.384**	-0.045	0.097	0.417
Hotel capacity	0.252**	0.206*	-0.972*	0.533	0.144
Child care facilities	-0.091	-0.088	0.054	0.570**	0.206
Land prices	0.121	0.175	-0.135	-0.266	0.571
East-West migration			5.048	-2.048	0.681
West-East migration			-3.495**	-2.708**	0.249
South-North migration			0.384	0.479	0.903
North-South migration			-0.095	0.257	0.787
Destination dummies[d]	No	No	Yes	Yes	
LL (Lower level)	-86646.9	-28762.9	-85348.5	-28369.7	
# of regional moves	31,465	10,225	31,465	10,225	
IIA fails[e](Small-Hsiao)	9/27	4/27	0/27	1/27	

Significance levels : † : 10% * : 5% ** : 1%

[a] Marginal effects and standard errors have been calculated as sample averages.
[b] LS: Less-skilled individuals with high-school degree or vocational training; HS: High-skilled individuals with tertiary education.
[c] P-values refer to test of difference between marginal effects for high- and less-skilled.
[d] Additional 27 destination dummies that are not shown, but available from the author upon request.
[e] Number of regions (out of 27) for which IIA fails at a significance level of 5%.

We can conclude that interregional economic differences affect the skill composition of interregional job flows for two main reasons. Firstly, higher wage levels disproportionately attract high-skilled migrants, especially high-skilled job-to-job changers, whereas interregional differences in wage dispersion, if at all, contribute only weakly to skill sorting across space. Secondly, unemployment differentials only exert a strong and significant effect on less-skilled job movers. Although the difference between both skill groups is not significant (p-value: 0.29), this finding is suggestive for the greater meaning of general job-finding conditions among less-skilled than among high-skilled job-movers.

Amenities and rents Compared to the impact of interregional income and job-finding differentials, amenity differentials as captured by the previously discussed proxies do not seem to have a strong impact on destination choices according to specification B in Table 5.5. To some extent, however, this may reflect that region-specific fixed effects soak up most amenity differentials. In a specification without such fixed effects, some amenity indicators such as the population size have a rather strong impact which may suggest that the relevance of such amenity factors may in fact be larger than suggested by specification B. On the other hand, specification A is likely to be biased due to unobserved regional heterogeneity. Specification B may thus be considered to give a less-biased, but also a rather conservative picture of the relevance of amenity differentials for the destination choices of job movers.

As regards different preferences for amenities between skill groups, parameter estimates for specification B in Table 5.5 are not contradictory to the idea that high-skilled individuals may have higher amenity valuations, but also do not present strong evidence in favour of this notion. Point estimates for the urban scale effect of higher population levels, for example, are indicative for higher valuations of consumer amenities among high-skilled job movers. Moreover, the availability of child care facilities as a specific type of public good significantly attracts only high-skilled job movers (+0.6pp). Table 5.6 suggests, however, that the already weak evidence in favour of higher amenity valuation among high-skilled movers vanishes when controlling for the type of job move. In particular, point estimates for the urban scale effect of higher population levels are twice as large for JJC than for UJC irrespective of skill level and are highly significant for JJC only. What is more, only job-to-job movers irrespective of skill level are significantly attracted to regions with a favourable child

care infrastructure. Although differences between JJC and UJC are not significant[19], these findings may weakly indicate that JJC have higher amenity valuations than UJC. Since JJC are relatively well-educated on average, such preferences may then also affect the skill sorting across space as weakly indicated in Table 5.5. Additional indicators such as regional crime rates or land prices do not significantly affect destination choices of job movers. Similarly, a higher population density seems to be a comparable disamenity for all sub-groups and thus also leaves the skill composition mainly unaffected.

Migration cost As expected, the likelihood of moving to a region significantly decreases with distance for all skill levels. Moreover, consistent with the theoretical framework, migration costs associated with migration distance are higher for less-skilled than for high-skilled job changers at a 10% significance level. In order to keep the probability of moving to region k constant if migration distance marginally increases from 100 to 101 km, the hourly wage level in k has to be 0.02 euros higher for high-skilled and 0.12 euros higher for less-skilled individuals.[20] Thus, the proportion of high-skilled following a particular migration path clearly increases with distance. According to Table 5.6, this finding is robust if the type of job move is controlled for.

For individuals born in West Germany, moving to the eastern part of the country is associated with a strong and significant disutility and thus additional migration costs while there is no additional utility assigned to the opposite direction for former East Germans. These costs may partially reflect economic disparities between both parts of Germany that are not captured by other covariates. Since no other migration paths yield any significant utility or disutility, however, covariates already seem to capture major regional disparities. Therefore, the disutility of moving to eastern Germany is likely to reflect some reluctance to cross the former border that is not explicable by observed regional disparities. Such reluctance has also been found by Büchel et al. (2002) in a study of migration intentions among West Germans. According to this study, only one third of those who are willing to change residential location are also willing to move to eastern Germany while more than

[19] In fact, for almost all parameters, establishing significant differences across skill groups turns out to be difficult due to imprecise estimates for at least one group.

[20] The change in wages that keeps the probability of moving to k constant if distance (km) increases is given by: $\frac{\partial wage}{\partial km} = \frac{\partial wage}{\partial log(km)} \frac{\partial log(km)}{km}$. Coefficient estimates are not shown, but are available from the author upon request.

50% are willing to leave the country. Thus, at least for individuals born in West Germany, the former border still seems to exist in their minds.

Table 5.6: Lower level marginal effects[a] $\frac{\partial P_{id|m}}{\partial z_{id}}$ by skill level and type of job mover in percentage points (pp), IAB-R01 1995-2001

	Less-skilled			**High-skilled**		
Variable	**JJC**	**UJC**	**p-value**[b]	**JJC**	**UJC**	**p-value**[b]
Median sector wage	0.62*	0.36	0.39	2.22**	1.42**	0.23
Wage variation	-0.24†	-0.49**	0.28	-0.06	0.35	0.32
Unemployment rate	-0.93**	-1.17*	0.56	-0.60	-0.60	0.95
Employment growth	0.49*	-0.15	0.15	0.24†	0.15	0.59
Share of HS employment	0.22	1.26**	0.21	-0.89	0.10	0.38
Log(Distance)	-6.07**	-6.25**	0.70	-4.22**	-4.68**	0.47
Population size	0.56**	0.23†	0.19	0.57**	0.22	0.37
Population density	-0.24*	-0.23†	0.78	-0.19†	-0.17	0.82
Crime Rate	0.02	-0.15	0.34	0.18	-0.20	0.30
Hotel capacity	-0.66*	-1.16**	0.16	0.37	0.95	0.48
Child care facilities	0.34†	-0.19	0.30	0.70*	0.23	0.35
Land prices	-0.08	-0.28	0.44	-0.28	-0.23	0.77
East-West migration	3.95	6.91	0.88	-2.18	-1.60	0.76
West-East migration	-3.31**	-3.63**	0.17	-2.77**	-2.43**	0.57
South-North migration	0.53	0.17	0.91	0.53	0.24	0.84
North-South migration	-0.12	-0.03	0.92	0.47	-0.39	0.84
Destination dummies[c]	Yes	Yes		Yes	Yes	
LL (Lower level)	-53615.7	-31591.5		-22410.0	-5930.1	
# of regional moves	19,906	11,559		8,093	2,132	
IIA fails[d](Small-Hsiao)	1/27	2/27		0/27	0/27	

Significance levels : † : 10% * : 5% ** : 1%

[a] Marginal effects and standard errors have been calculated as sample averages. See previous section for details.

[b] P-values refer to test of difference between marginal effects for high and less-skilled.

[c] Additional 27 destination dummies that are not shown, but available from the author upon request.

[d] Number of regions (out of 27) for which IIA fails at a significance level of 5%.

Upper level estimates Table 5.7 shows marginal effects on the marginal probability of leaving the local region, i.e. the probability of experiencing an interregional instead of an intraregional job move. The estimates include the inclusive value estimate $\hat{iv}_{im}$ from the lower level specification B. This inclusive parameter reflects the expected utility that an individual derives from migration. The corresponding parameter estimate ζ indicates whether pull factors are important in determining mobility decisions. According to a test of equal parameters across sub-groups, high-skilled job-to-job changers are significantly more responsive to pull factors than other sub-groups. As a consequence, the share of interregional movers who are high-skilled slightly increases if other labour markets gain in attractiveness. Apart from the inclusive value, there are a number of additional upper level covariates that significantly affect the decision to change a job interregionally. Across all sub-groups, younger, better skilled and previously well-earning job changers are more likely to be interregionally mobile. The latter two findings may both reflect higher migration propensities among high-ability individuals since the previous wage income is likely to capture some heterogeneity in ability that is unexplained by formal education. Among the employment history indicators, having previously been recalled dramatically reduces the likelihood of changing a job interregionally because these individuals tend to be recalled locally again and may simply not be looking for jobs elsewhere. Longer average tenure also reduces the probability of leaving the local region, probably due to the regional attachment that comes with a long job tenure. Furthermore, migration levels increased during the observation period from 1995 to 2001. This is in line with Heiland (2004) who finds that increasing migration levels coincided with a period of stagnation in eastern Germany in the mid to late 1990s. Finally, the estimates suggest a much higher probability of changing a job interregionally for less-skilled East Germans as compared to West Germans. This may mainly reflect unfavourable employment conditions that force especially less-skilled individuals in eastern Germany to look for jobs in alternative locations.

Table 5.7: Upper level marginal effects[a] $\frac{\partial P_{im}}{\partial w_i}$ for specification B by skill level and type of job move (in pp), IAB-R01 1995-2001

	JJC & UJC		**JJC**		**UJC**	
Covariates	**LS**	**HS**	**LS**	**HS**	**LS**	**HS**
Age 25-30	0.86*	5.15**	1.23**	4.61**	0.27	5.19**
Age 30-35	1.08**	4.15**	1.29**	3.37**	0.80*	6.81**
Age 40-45	-0.07	-3.90**	-0.11	-4.62**	-0.07	-2.08
UJC	-1.92**	4.66**	n/a	n/a	n/a	n/a
Unskilled	-2.07**	n/a	-2.32**	n/a	-1.51*	n/a
Born in East Germany	9.27*	-3.26	9.08†	-3.36	9.26*	-2.60
2nd wage quintile	0.17	-6.27**	0.01	-7.44**	0.25	-2.48
3rd wage quintile	2.02**	-2.02*	2.69**	-3.31**	1.29	2.13
4th wage quintile	5.70**	3.87**	6.41**	2.70†	4.27**	7.54**
5th wage quintile	12.9**	9.51**	13.66**	7.58**	14.6**	18.1**
Average tenure	-0.75**	-0.87**	-0.78**	-1.05**	-0.67**	0.22
Mth. non-employed	-0.13	-0.50*	-0.28**	-0.37	-0.05	-0.70**
Prev. recall	-16.41**	-20.70**	-3.17**	-6.66*	-16.3**	-35.5**
Multiple job changes	1.03**	-1.07	1.92**	-0.77	-0.09	-2.16
1996	-1.00*	-0.45	-0.44	1.09	-1.18*	-5.30
1997	-0.43	1.88	-0.18	2.74*	-0.16	-0.48
1998	0.55	2.38*	0.41	2.68**	1.19†	2.21
1999	0.44	2.69†	0.85	2.93	0.45	2.66
2000	2.06**	1.82	2.24*	1.52	2.49**	5.14*
2001	1.86**	3.29*	2.35**	3.22†	1.93**	5.17†
Other covariates[b]	X	X	X	X	X	X
ζ_m[c]	0.31**	0.47**	0.28**	0.46**	0.34**	0.29**
LL (upper level)	-77495.2	-16892.4	-47106.2	-13473.5	-29812.4	-3335.4
# of job moves	175,587	26,457	95,938	21,040	79,649	5,417

Significance levels : † : 10% * : 5% ** : 1%

[a] Marginal effects and standard errors have been calculated as sample averages. See previous section for details.

[b] Includes 13 sector of activity dummies, 9 types of occupation dummies, 27 origin dummies. Full estimation results are available from the author upon request.

[c] Displays coefficient estimate instead of marginal effect.

Simulation Results Based on the preceding estimation results, this section simulates how the level and skill composition of job matching flows changes in a scenario of economic convergence between western and eastern Germany. This is an interesting case to study because of the continued regional employment and wage disparities between both parts of Germany and the resulting loss of population and human capital in eastern Germany that has been discussed in section 5.4. As a drawback of the job-changing framework, however, any such simulation can only capture the changing spatial job matching pattern of those who move jobs, but abstracts from induced job mobility. Still, looking at the spatial job matching pattern of job movers should yield some insights into which factors may, for example, help in attracting human capital to eastern Germany. I therefore simulate job matching patterns for a scenario of economic convergence based on specification B in Table 5.5. I simulate mobility patterns by using the observed wage level, wage variation, unemployment rate and employment growth for all western regions while adjusting the corresponding values for eastern regions according to the following formula:

$$\tilde{z_e^s} = \tilde{z_e} + (\frac{1}{N_w} \sum_{k \in w} \tilde{z_w} - \frac{1}{N_e} \sum_{k \in e} \tilde{z_e}) \tag{5.8}$$

where $\tilde{z_e^s}$ refers to the simulated standardised value[21] for the eastern region and N_e (N_w) denotes the number of eastern (western) regions. This simulation results in higher wage levels, increasing wage variation, lower unemployment rates, increasing employment growth for high-skilled and slightly increasing employment growth for less-skilled individuals in eastern Germany. Moreover, this mean convergence maintains regional disparities within eastern Germany. Appendix F shows average observed and simulated values for eastern Germany.

Table 5.8 shows the predicted mobility pattern based on specification B in Table 5.5 for less-skilled and high-skilled individuals. Note that the predicted mobility pattern strongly resembles the observed pattern in Table 5.2 which suggests some explanatory power of the econometric model. Table 5.9 shows percentage point differences between the predicted and the simulated mobility pattern in case of an economic convergence. Due to much higher responsiveness to pull factors among high-skilled job-to-job movers than among other sub-

[21] As discussed in section 5.5, lower level covariates z_{ik} refer to the difference of standardised values. The simulated standardised $\tilde{z_e^s}$ are thus used to calculate the differences in case of a simulated economic convergence.

groups, the probability of leaving the local region strongly increases for high-skilled individuals in western states and strongly decreases in eastern states compared to much weaker reactions for their less-skilled counterparts. More importantly, economic convergence attracts job movers of all skill levels and from all regions to eastern Germany. Pull factors are again much stronger for high-skilled than for less-skilled individuals however. In fact, the probability of moving to the eastern states more than triples for high-skilled individuals, but less than doubles for less-skilled individuals.

Table 5.8: Predicted job matching pattern by skill level, IAB-R01, 1995-2001

			Destination (in %)				
Origin	**Skill level**	**Obs.**	**Stay Home**	**East**	**North**	**Mid**	**South**
East	Less-skilled	49,935	84.0	8.2	2.6	2.6	2.6
	High-skilled	4,862	69.1	14.8	4.6	6.2	5.4
North	Less-skilled	28,009	82.1	2.7	7.5	5.9	1.8
	High-skilled	3,913	58.8	5.1	12.8	16.0	7.4
Mid	Less-skilled	56,085	79.5	1.2	2.8	12.5	4.0
	High-skilled	10,364	58.4	2.3	5.5	22.9	10.8
South	Less-skilled	41,558	83.3	1.5	1.0	4.4	9.8
	High-skilled	7,318	62.1	2.5	3.4	13.2	18.8

As a consequence, economic convergence affects net job flows between both parts of Germany and changes the skill composition of west-east and east-west flows as can be seen in Table 5.10. Besides looking at the effects of a full economic convergence as described above, Table 5.10 also identifies the main sources of the simulated change by looking at the effects in case of an isolated convergence of wage levels, wage dispersion, unemployment rates and employment growth, respectively. As suggested by the previous estimation results, the increasing skill level of west-east flows from 23.5% to 39.6% in case of a full economic convergence is mainly driven by increasing wage levels in eastern states. Higher wage inequality in eastern regions also increases the skill level of west-east flows. This is due, however, to an increasing net outflow of less-skilled job movers. By contrast, converging wage levels not only strongly increase the share of high-skilled west-east migrants, but also substantially raise net migration as has also been suggested by Burda and Hunt (2001).

Table 5.9: Simulated change in the spatial pattern of job movements by skill level in case of an economic convergence between western and eastern Germany, IAB-R01 1995-2001

			Destination (pp change)				
Origin	**Skill level**	**Obs.**	**Stay Home**	**East**	**North**	**Middle**	**South**
East	Less-skilled	49,935	1.38	2.15	-1.18	-1.18	-1.18
	High-skilled	4,862	5.27	6.41	-3.31	-4.47	-3.91
North	Less-skilled	28,009	-0.85	2.58	-0.92	-0.61	-0.21
	High-skilled	3,913	-4.97	14.92	-3.61	-4.23	-2.10
Mid	Less-skilled	56,085	-0.39	1.42	-0.17	-0.62	-0.24
	High-skilled	10,364	-2.55	8.69	-0.90	-3.49	-1.75
South	Less-skilled	41,558	-0.47	1.58	-0.08	-0.31	-0.72
	High-skilled	7,318	-2.86	9.03	-0.63	-2.27	-3.27

In case of full convergence, it is mainly lower unemployment levels that further raises the number of net migrants, mainly due to an increased net migration of less-skilled job changers. Thus, while higher wage levels turn out to be an effective means of attracting human capital to eastern Germany, the net outflow from eastern to western regions can only be reversed by a combination of higher wage levels and lower unemployment rates.

Table 5.10: Net job flows and induced net employment change by skill level and share of high-skilled migrants for various scenarios, IAB-R01 1995-2001

	Net migration		**Net emp. change**		**Share HS migrants**	
	LS	**HS**	**LS**	**HS**	**east-west**	**west-east**
Observed	-1,447	-183	-1.4 %	-1.2 %	16.5 %	18.7 %
Predicted	-1,847	-154	-1.8 %	-1.0 %	16.7 %	23.5 %
Isolated convergence of						
Wage level	-376	1,675	-0.4 %	10.7 %	8.9 %	41.9 %
Wage variance	-2,812	-134	-2.7 %	-0.9 %	14.7 %	27.8 %
Unemployment rate	1,505	247	1.5 %	1.6 %	21.1 %	18.5 %
Employment growth	-1,803	-99	-1.7 %	-0.6 %	16.4 %	24.2 %
Full convergence	2,096	2,559	2.0 %	16.4 %	9.2 %	39.6 %

Note: Employees by skill level are computed based on the IAB-R01 at the beginning of the observation period (01/01/1995).

5.7 Conclusion

This paper has identified major determinants of the skill composition of internal job matching flows in Germany by looking at destination choice patterns of heterogeneous job movers. Since regional economic prospects critically hinge on the skill composition of internal migration flows, the analysis provides some insights into how policy may contribute to convergence and integration. Such insights are of particular value in light of the continued brain drain from eastern to western Germany. As an extension to previous studies concerning the relationship between destination choices and skill level, this study has examined whether different destination choice patterns of job-to-job changers and job changers after unemployment contribute to skill-sorting across space. Moreover, the analysis takes account of unobserved regional heterogeneity which proved important to avoid biases arising from the omission of unobserved regional characteristics. Using a partially degenerate nested logit analysis, this paper comes to the following main conclusions:

- Interregional income differentials affect the skill composition of job matching flows mainly because high-skilled job movers are much more responsive to interregional variation in the wage level than their less-skilled counterparts. By contrast, there is only weak evidence that wage inequality induces some skill sorting as shown for the US by Hunt and Mueller (2004). This comparatively weak finding in favour of the extended Roy selection model may reflect that central wage bargaining in Germany leaves little scope for local wage agreements.

- Interregional unemployment differentials only exert a significant effect on less-skilled job seekers. This finding is suggestive for a greater meaning of general job-finding conditions among less-skilled than among high-skilled job movers that may affect the skill composition of job matching flows.

- Amenity valuations only seem to contribute weakly to skill sorting across space. Moreover, if at all, skill sorting seems to be induced by higher amenity valuations of job-to-job changers compared to job changers after unemployment. There is no evidence in favour of higher amenity valuations among high-skilled individuals when controlling for the type of job move.

- High-skilled job movers face lower migration costs such that the proportion of high-skilled migrants strongly increases with migration distance.

- High-skilled job-to-job movers are more responsive to pull factors than all other sub-groups. Improving destination conditions thus disproportionately mobilise this group which affects the skill composition of internal job matching flows.

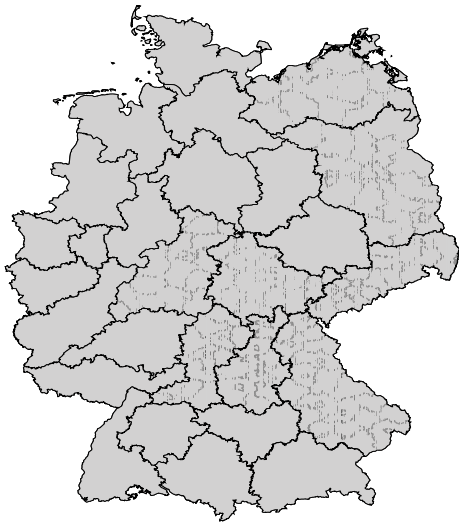
These findings imply that rising wage levels in eastern Germany during the 1990s have been an effective means of preventing a stronger brain drain. However, wages have risen at the cost of higher unemployment levels which mainly boosted east-west migration of less-skilled individuals. A simulated economic convergence between eastern and western Germany shows that higher wage levels are the most effective means of attracting human capital to eastern Germany, but that the net loss of population can only be reversed by lower unemployment rates. If maintaining the future viability of eastern Germany is a pronounced policy objective, the findings in this paper thus advocate policies that foster wage convergence without further increasing eastern unemployment levels. For this purpose, policy measures that help in closing the productivity gap between eastern and western Germany may be a first choice.

Finally, the study points to a number of upcoming research tasks. First of all, although specification tests suggest that the nested logit approach is well specified, a less restrictive specification such as a mixed logit may be used as an alternative approach in future research. Secondly, the job-matching framework of the analysis does not allow to endogenously model the job mobility decision. Thus, the paper explains the spatial distribution of job matches conditional on changing a job without taking into account that certain labour market conditions may induce or retard job mobility. Moreover, to the extent that labour market conditions affect the composition of job movers with respect to the type of job change, for example, spatial job matching patterns and thus also the interregional competition for jobs may differ across the business cycle. Extending the analysis to endogenously model the job mobility decision thus is an important future research direction. Finally, due to data restrictions, the analysis leaves out highly mobile and important labour market segments, namely single females and first time job entrants such as university graduates. Future research should examine destination choice patterns of these segments because they strongly affect the skill composition of internal job matching flows.

Appendix

A – The definition of regional boundaries for the analysis

Regional boundaries for the destination choice analysis are based on the 97 planning regions. These regional entities have been delineated by the Federal Office for Building and Regional Planning (*Bundesamt für Raumordnung und Bauwesen*) according to commuting ranges such that the majority of commutes occur within a planning region. For a feasible estimation of the destination choice model, the number of destination regions had to be reduced. For this purpose, an algorithm was used that lumps together planning regions by minimising external commuting linkages between adjacent planning regions subject to the constraint that no more than five regions may be lumped together and that western and eastern regions remain separated. The latter restriction ensures that flows between western and eastern Germany can still be identified. External commuting linkages between the planning regions have been provided by the Federal Office for Building and Regional Planning for the year 2003. Although commuting linkages may change over time, I assume that linkages from 2003 are still quite similar to the relevant commuting linkages during the observation period 1995 to 2001. Using the described aggregation algorithm results in the following 27 aggregated planning districts:



B – Definition and data sources of lower-level covariates

Variable	Definition	1 yr Lag	Data Source[a]
Covariates with area and individual variation			
Median sector wage	Median wage in individual i's sector of activity l ($l = 1..13$)	No	A
Sector employment growth[b]	Biennial employment growth in individual i's skill group	Yes	B
Covariates with area variation			
Wage variance index	Regional wage percentile ratio divided by aggregate percentile ratio	No	A
Unemployment rate	Average yearly unemployment rate	Yes	C
Share HS employment	Share of high-skilled employment	Yes	B
Log(Distance)	Log of average distance between all county capitals of any two regions	-	D
Population size	Number of residents in 100,000	Yes	E
Population density	Number of residents (in 100) per km^2	Yes	E
Crime Rate	Total offenses per 100 residents	No	F
Hotel capacity	Number of hotel beds per 1000 residents	No	E
Child care facilities	Places in day care for children and youth per 1,000 residents	No	E
Land prices	Land prices in 100 euros per m^2	No	E

[a] A - Own calculation based on IAB-R01 1995-2001. See Appendix C for details on the calculation.
B - Own calculation based on IAB-R01 1993-2001
C - Federal Employment Agency (*Bundesagentur für Arbeit*)
D - Own calculations based on the grid position of county capitals
E - Federal Statistical Bureau (*Statistisches Bundesamt*)
F - European Regional Crime Database, Entorf and Spengler (2004)

[b] I distinguish between employment growth for high-skilled individuals with a tertiary education and less-skilled individuals.

C – Estimating moments of the regional wage distribution

The observed regional wage distribution reflects interregional differences in both skill prices and skill mix. For this reason, Hunt and Mueller (2002) control for interregional differences in the skill mix by estimating key parameters of a standardised regional wage distribution that is comparable across regions. Mainly following their methodology, I separately estimate Mincerian-type wage equations for each region k based on the IAB-R01. Since wages in the IAB-R01 are top-coded at the income level above which there is no obligation to be socially insured, I estimate a tobit model. I restrict the estimation to prime age males who are full-time employed on January 1st and include educational attainment, experience, occupation

type and sector of activity as covariates. I predict the standardised wage distribution for region k by using the entire sample of prime age males and the coefficient estimates for region k. Since the same sample is applied to each region, this procedure controls for interregional differences in skills and experience levels and thus yields a standardised wage distribution. Due to the censoring in the data, it is not possible to consistently estimate the moments of this standardised distribution. For this reason, I use percentiles which are unaffected by the censoring of the data to appropriately measure interregional wage differences. Therefore, I estimate the median wage by sector of activity as an indicator of interregional differences in the sector-specific wage levels. As an indicator of the wage variance in region k, I calculate the difference between the 80th and 20th wage percentile based on the standardised wage distribution for region k and divide it by the corresponding percentile ratio of the wage distribution when pooling all regions. If this wage variance index is larger than 1, the wage inequality in region k exceeds the average wage inequality.

D – Sample averages for lower level covariates by sub-sample

	JJC		**UJC**	
Covariates	**LS**	**HS**	**LS**	**HS**
Median sector wage	0.104	0.055	0.139	0.110
Wage variance index	0.066	0.013	0.085	0.034
Unemployment rate	-0.114	-0.057	-0.146	-0.137
Employment growth	0.026	0.028	0.038	0.043
Share of HS employment	-0.009	0.028	-0.009	0.125
Log(Distance)	5.163	5.307	5.209	5.276
Population size	0.012	0.014	0.022	0.050
Population density	0.014	0.003	0.041	0.032
Crime Rate	-0.051	-0.020	-0.066	-0.060
Hotel capacity	0.010	0.002	-0.019	0.012
Child care facilities	-0.048	0.006	-0.062	-0.050
Land prices	0.094	0.071	0.122	0.205
# of interregional moves	19,906	8,093	11,559	2,132

* Except for log(distance), all covariates refer to the difference between the standardised value for the destination (d) and the origin (o) region. Thus a value of 1 indicates a difference of one standard deviation between d and o.

** JJC - Job-to-job changer; UJC - Job changers after unemployment; LS - Less-skilled; HS - High-skilled

E – Sample averages for upper level covariates by sub-sample

	JJC		UJC	
Covariates	**LS**	**HS**	**LS**	**HS**
Migrant	0.21	0.38	0.15	0.39
Age (Reference: Age 30-35)				
25-30	0.29	0.15	0.27	0.13
30-35	0.29	0.37	0.27	0.32
40-45	0.19	0.20	0.22	0.26
Wage quintile in previous job[a](Reference: 1st wage quintile)				
2nd	0.25	0.08	0.31	0.18
3rd	0.16	0.09	0.17	0.15
4th	0.10	0.19	0.07	0.17
5th	0.07	0.39	0.02	0.19
Employment history and other covariates				
Born in East Germany	0.21	0.13	0.33	0.22
Multiple job changes[b]	0.73	0.71	0.85	0.73
Prev. average tenure (yrs.)	2.91	2.44	1.74	1.76
Months prev. non-employed	1.04	0.76	2.24	1.97
Prev. recall by employer	0.01	0.01	0.18	0.03
Previous sector of activity (Reference: Agriculture and Fishing)				
Primary industry	0.06	0.05	0.06	0.04
Invest. goods/engineering	0.08	0.08	0.05	0.07
Invest. goods/vehicles	0.07	0.11	0.04	0.07
Cons. goods/ food process.	0.07	0.04	0.07	0.04
Construction	0.17	0.05	0.37	0.10
Wholesale trade	0.08	0.07	0.05	0.07
Retail	0.07	0.03	0.05	0.04
Transport/Communication	0.10	0.03	0.06	0.03
Financial services	0.17	0.32	0.09	0.22
Domestic services	0.05	0.02	0.04	0.03
Social services	0.04	0.15	0.05	0.22
Public authorities	0.01	0.02	0.02	0.04
Previous type of occupation (Reference: Agricultural work)				
Blue-collar work	0.51	0.05	0.66	0.12
Salesmen	0.07	0.06	0.04	0.06
Technical work	0.06	0.35	0.03	0.29
Clerical work	0.06	0.11	0.03	0.10
White-collar work	0.05	0.26	0.02	0.16
Health-related/Teaching/Consulting	0.02	0.12	0.01	0.18
Other service jobs	0.20	0.05	0.16	0.07
Year of job move (Reference: 1995)				
1996	0.13	0.11	0.15	0.14
1997	0.13	0.12	0.17	0.15
1998	0.13	0.14	0.15	0.15
1999	0.15	0.17	0.15	0.15
2000	0.16	0.18	0.13	0.13
2001	0.15	0.18	0.13	0.13
# of job moves	95,935	21,040	79,649	5,417

[a] Wage quintile of the wage distribution of all full time employees observed on January 1st of each year (Data: IAB-R01).
[c] Indicator whether an individual contributes two or more observations (i.e. job changes) to the sample.
[d] The duration of the previous spell refers to the previous job tenure for JJC and to the unemployment period for UJC.

F – Average observed and simulated unemployment rate, wage level, wage dispersion and employment growth for eastern Germany

Indicator	$\frac{1}{N_e}\sum_{k\in e}\tilde{z}_e^s$	$\frac{1}{N_e}\sum_{k\in e}\tilde{z}_e$
Median wage in agriculture	0.50	-1.75
Median wage in primary ind.	0.49	-1.73
Median wage in inv. good/engineering	0.49	-1.70
Median wage in inv. goods ind./vehicles	0.48	-1.69
Median wage in cons. goods/food process.	0.50	-1.74
Median wage in construction	0.52	-1.80
Median wage in wholesale trade	0.48	-1.68
Median wage in retail	0.47	-1.70
Median wage in transport/communication	0.48	-1.69
Median wage in financial services	0.50	-1.76
Median wage in domestic services	0.42	-1.47
Median wage in social services	0.50	-1.76
Median wage in public authorities	0.47	-1.66
Wage variance	0.46	-1.62
Unemployment rate	-0.46	1.60
Emp. growth for less-skilled jobs	-0.39	-0.56
Emp. growth for high-skilled jobs	1.02	0.25

Note: Average simulated values for eastern regions correspond to the observed average values for western regions, i.e. $\frac{1}{N_e}\sum_{k\in e}\tilde{z}_e^s = \frac{1}{N_w}\sum_{k\in w}\tilde{z}_w$.

Concluding Remarks and Outlook

Levels of internal migration in Germany are relatively low in international comparison and may contribute to lower overall employment levels, lower economic growth and persistent regional employment disparities. Increasing levels of internal mobility may thus be a means of realising potential welfare gains. Against this background, the main objective of this thesis was to shed light on the determinants of mobility for heterogeneous labour market segments and to identify the scope for policy makers to raise mobility levels. In particular, the thesis focused on mobility decisions of unemployed jobseekers because the willingness and ability of unemployed individuals to be interregionally mobile should be of particular importance if geographic mobility is to contribute to higher overall employment levels and to an accelerated regional convergence. By looking at unemployment experiences and mobility decisions of unemployed jobseekers, this dissertation provides a number of interesting insights concerning the responsiveness of unemployed jobseekers to regional labour market conditions and the extent to which labour market institutions such as the unemployment compensation system create obstacles to the mobility of unemployed individuals. While the details of the individual results in each preceding paper is to be found in the corresponding conclusions, there are a number of main findings that echo through all of these papers and that are highly relevant from a policy perspective.

First of all, the findings from paper (2) and (3) indicate that unemployment experiences and the geographic mobility of unemployed individuals are clearly dominated by individual characteristics such as education, age, but also the employment history. In addition, there is some heterogeneity among unemployed jobseekers in the responsiveness to regional labour market conditions. While skilled men and well-earning singles respond to a relatively weak local labour demand by higher migration probabilities, unskilled men, low-earning individuals and women have been found to stay in regions despite unfavourable job-finding prospects. These latter groups thus constitute a rather immobile labour market segment that is partic-

ularly dependent on local labour market conditions. A regional labour demand shock thus prolongs unemployment and builds up an increasing level of regional long-term unemployment, especially among these immobile labour market segments. Hence, the contribution of geographic mobility to regional convergence is likely to be rather limited which is in line with findings by Möller (1995) that adjustment processes after region-specific shocks tend to be slow in Germany. Moreover, paper (3) indicates that many of these rather immobile individuals are likely to end up in subsidised employment. Since leaving the labour force is likely to be another important option for many immobile individuals, modelling transitions out of the labour force would certainly be an interesting extension, but the available administrative data unfortunately does not allow for observing such transitions. If such data became available, we could examine at the micro level what macro-oriented studies have already pointed at, namely that a region-specific shock often increases the non-participation among the local labour force (Decressin and Fatàs, 1995).

Papers (2) to (4) provide some evidence that labour market institutions[22] may at least provide a partial explanation for low mobility levels and the weak responsiveness to regional labour market conditions among many unemployed jobseekers. In particular, passive labour market policies appear to have an impact on the duration of unemployment and the probability of leaving unemployment via migration. For all groups of unemployed individuals, papers (2) and (3) indicate that being entitled to receive unemployment benefits (UB) for an extended period is associated with lower migration probabilities. Since these findings need not be interpretable as a pure causal effect of UB receipt, paper (4) re-examined these findings by exploiting the variation in entitlement length from a labour market reform in 1997. The results from paper (4) are again suggestive for the mobility-reducing effect of unemployment benefits, at least among high-skilled individuals for whom the shortening of UB receipt entails the strongest income loss. The missing reaction to the 1997 reform among low-skilled individuals, however, cannot be interpreted as a missing evidence for the impact of the unemployment compensation system on the unemployment experiences of this group. This is because individuals with low pre-unemployment earnings often receive social benefits in addition to unemployment compensation and thus reach income replacement rates up to

[22]Since the empirical results in this thesis have all been obtained in a period prior to the latest Hartz reforms, labour market policies refer to the previous institutional setting.

or even exceeding 100% that are independent of whether receiving unemployment benefits or unemployment assistance. The 1997 reform that cut the length of UB receipt should thus have much weaker effects on low-earning individuals despite strong disincentives to seek employment and be interregionally mobile. The evidence from this natural experiment in favour of a mobility-reducing impact of unemployment benefits among those for whom the threat of entitlement loss after the exhaustion of UB entitlements is likely to be strong suggests that the barrier to mobility caused by the unemployment compensation and welfare system may actually be quite severe for individuals with extensive income replacement rates. The findings in papers (2) and (3) concerning an extremely long duration of unemployment and extremely low migration rates among groups of unemployed who are likely to receive increased income replacement rates are consistent with this notion. The available administrative data does not contain sufficient information to identify additional welfare receipt and actual income replacement rates. This points towards the need to re-examine the link between the level of basic income support and mobility if such data became available in the future.

Limited disincentives to mobility also stem from active labour market programs (ALMP). At least for women and married men, papers (2) and (3) present some evidence in favour of a limited regional locking-in effect in regions with an extensive provision of active labour market programs. Apparently, active labour market programs are a substitute for leaving the region that distracts rather immobile groups of jobseekers from job search in the regular labour market. The research findings thus indicate a limited scope for increasing mobility by a reduction of ALMP. On the other hand, the effect of participating in ALMP on mobility has been found to depend on the type of program and may be positive for certain training programs (Lindgren and Westerlund, 2003). A concluding assessment to what extent a reduction of ALMP may contribute to higher internal migration in Germany thus calls for additional research on the effects of participating in ALMP on geographic mobility in order to complement the findings from this thesis.

The thesis has thus identified some scope for policy makers to increase mobility among unemployed jobseekers. In particular, the findings indicate that a reduction of transfer receipt in the case of unemployment is likely to foster migration and to shorten the duration of unemployment among unemployed jobseekers. Since a certain minimum level of income

support in case of unemployment is both legally defined and societally desired, however, there are limits to the reduction of unemployment compensation as a means of promoting geographical mobility. In particular, a reduction of the basic income support is politically conceivable only in combination with in-work welfare programs such as the negative income tax credit in the US or the working tax credit in the UK (see e.g. Kaltenborn and Pilz, 2002). Alternatively, workfare programs that require individuals to work to be eligible for income support might also improve not only work incentives, but also the willingness to be interregionally mobile. The effect of such programs on the geographical mobility of unemployed jobseekers, however, is vastly understudied which points towards future research needs. Moreover, the likely limits to a reduction of basic income support also highlight the need to take account of alternative ways how to promote mobility among unemployed individuals but also among the employed workforce. Additional barriers to mobility are, for example, likely to also stem from housing policies. Social housing policies as well as high transaction costs incurred when selling and buying real estate have been found to reduce geographic mobility among its tenants in Germany (Barcelo, 2003). Research on the impact of such housing policies may thus help in identifying additional scope for increasing internal migration in Germany.

Although removing obstacles to geographic mobility is an important policy issue in order to realise the potential welfare gains from a higher level of geographic mobility, there are a number of possible downsides of a higher level of geographic mobility. Welfare losses in the form of adjustment costs, a loss of social capital that may even result in reduced fertility, an increasing social and spatial segregation if weaker social groups are displaced from certain areas, or a possible divergence of the regional system due to the selective nature of migration call for some compensating social and economic policies to mitigate the possible negative consequences of higher mobility levels. Although a comprehensive assessment of the negative consequences of geographic mobility and possibilities to cushion such effects clearly goes beyond this thesis, another contribution of this dissertation was to provide insights on how to cushion one particular downside of a higher level of geographic mobility, namely its potentially divergent effect on the regional system.

For this purpose, the last part of this dissertation looked at the destination choices of different skill groups because the composition of migration flows with respect to productivity-

related characteristics may affect whether geographic mobility rather fosters convergence or divergence. The last paper thus examined the mechanisms that result in a sorting of skill groups across space and suggests that the composition of migration flows to some extent can be shaped by policy. In particular, paper (5) indicates that a continued brain drain from eastern to western Germany that is likely to reinforce the existing employment and wage disparities between both parts of the country may be mitigated or even reversed by means of increasing wage levels in eastern Germany. However, the findings from this paper also suggest that increasing wage levels in eastern Germany in order to attract high-skilled individuals may likely boost west-east migration of less-skilled workers if higher wages widen the gap between productivity and wages and thus cause rising unemployment. In addition to raising regional productivity levels, one way of increasing wages for high-skilled individuals in eastern Germany without widening the gap between wages and productivity for less-skilled labour market segments thus is a regionally tailored level of wage inequality. The current system of collective wage bargaining in Germany, however, limits the scope for regionally tailored wage agreements (OECD, 2005). An insufficient wage flexibility at the regional level may thus limit the scope for attracting human capital to eastern Germany. In addition, a resulting gap between productivity and wage levels in depressed regions also limits the stimulation of new labour demand and thus contributes to the persistent nature of east-west unemployment disparities in Germany (de Koning et al., 2004).[23] Removing barriers to the mobility of labour demand by increasing the scope for regionally tailored wage agreements may thus be an important supplement to policies that remove barriers to labour supply. Such a policy approach may help in raising mobility levels while at the same time counteracting a possible divergence of the regional system. In an increasingly flexible and mobile world, the relevance of a comprehensive policy approach that also considers the possible downsides of geographic mobility and takes account of the heterogeneous incentives and disincentives that determine individual mobility behaviour cannot be overemphasised. As its major contribution, this dissertation has identified some scope for designing such policies.

[23] A similar mechanism has been considered to lead to persistent regional employment disparities between the north and the south of Italy (Brunello et al., 2001).

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Berufserfahrung und Praktika

Wissenschaftliche Mitarbeiterin am Zentrum für Europäische Wirtschaftsforschung (ZEW) in Mannheim	**seit 10.2002**
Tutorin für Statistik im Fachbereich Geographie an der Universität Bonn	**11.00-04.02**
Studentische Hilfskraft bei Prof. Dr. M.J.M. Neumann am Institut für Internationale Wirtschaftspolitik der Universität Bonn	**03.01-07.01**
Praktikum bei der Unternehmensberatung McKinsey & Company	**07.-10.00**
Wissenschaftliche Mitarbeiterin an einem Forschungsprojekt der Weltbank zu den Anpassungsprozessen am ägyptischen Arbeitmarkt der 1990er Jahre (Leitung: PhD Ragui Assaad, University of Minnesota, MN, USA)	**05.-07.00**
Studentische Hilfskraft bei PhD Ragui Assaad im Fachbereich Public Affairs an der University of Minnesota, MN, USA	**01.-05.00**
Studentische Hilfskraft bei Prof. Dr. H.- D. Laux im Rahmen eines Forschungsprojektes zur Integration von Einwanderern in den USA	**04.98-07.99**
Praktikum bei der Stadtentwicklung der Stadt Köln	**07.97-08.97**

Schulischer und akademischer Werdegang

Promotion zum Dr.rer.pol. im Fachbereich Wirtschaftwissenschaften der TU Darmstadt, Titel der Dissertation: *The Geographic Mobility of Heterogeneous Labour in Germany*	**05.07**
Zweitstudium der Volkswirtschaftslehre an der Universität Bonn Vordiplom mit mathematischem Schwerpunkt	**04.98-10.01**

Studium der Geographie an der Universität Bonn Abschluss als Diplom-Geografin Diplomarbeit: *Determinants of Female Participation in Wage Employment under Constrained Geographical Mobility. The Case of Egypt*	**10.96-05.02**
Abitur am Landrat-Lucas-Gymnasium in Leverkusen	**06.96**

Stipendien

Fulbright-Stipendium für ein Studienjahr an der University of Minnesota, MN (USA) im Fachbereich Wirtschaftsgeographie	**07.99-07.00**
Stipendiatin der Studienstiftung des Deutschen Volkes	

Wissenschaftliche Vorträge (Auswahl)

• European Society for Population Economics (ESPE) 21th Annual Conference, Chicago • 2. Arbeitstreffen des AGF (Anglo-German Foundation) Projekts „Creating Sustainable Growth in Europe",Lezignan-Corbieres • 4. Arbeitstreffen des DFG-Schwerpunktprogramms „Flexibilisierungspotenziale bei heterogenen Arbeitsmärkten", Mannheim	**2007**
• Jahrestagung des Vereins für Socialpolitik, Bayreuth • Annual Meeting of European Association of Labour Economists, Prag • Workshop; „Labour Market Flexibility, Inter-firm and Inter-regional Mobility", Regensburg • 2. Arbeitstreffen des AGF (Anglo-German Foundation) Projektes „Creating Sustainable Growth in Europe", London	**2006**

- Jahrestagung des Vereins für Socialpolitik, Bonn **2005**
- IAB-Nutzerkonferenz, Nürnberg
- SOLE/EALE 2nd World Conference, San Francisco
- Fakultätsseminar, TU Darmstadt
- Arbeitstreffen des DFG-Schwerpunktprogramms „Flexibilisierungspotenziale bei heterogenen Arbeitsmärkten", Mannheim

- Interdisciplinary spatial statistics workshop, Paris **2004**

Beiträge in referierten Fachzeitschriften

Arntz, Melanie und Ralf Wilke (2007), *Unemployment Duration in Germany: Individual and Regional Determinants of Local Job Finding, Migration and Subsidized Employment*, Regional Studies (im Erscheinen).

Arntz, Melanie und Ralf Wilke (2007), An application of cartographic area interpolation to German administrative data, *Allgemeines Statistisches Archiv* (im Erscheinen).

Arntz, Melanie, Stefan Boeters und Nicole Gürtzgen (2006), Alternative Approaches to Discrete Working Time Choice in an AGE Framework, *Economic Modelling* 23, 1008-1032.

Arntz, Melanie, Jochen Michaelis und Alexander Spermann (2006), Reforming Long-term Care in Germany: Preliminary Findings from a Social Experiment with Matching Transfers, *Swiss Journal of Economics and Statistics*, 37-42.

Assaad, R. und Melanie Arntz (2005), Constrained Geographical Mobility and Gendered Labor Market Outcomes Under Structural Adjustment: Evidence from Egypt, *World Development* 33 (3), 431-454.

Arntz, Melanie und Alexander Spermann (2005), Soziale Experimente mit dem Pflegebudget (2004-2008) - Konzeption des Evaluationsdesigns, *Sozialer Fortschritt* 54(8), 181-191.

Arntz, Melanie und Alexander Spermann (2004), Wie lässt sich die gesetzliche Pflegeversicherung mit Hilfe personengebundener Budgets reformieren?, *Sozialer Fortschritt* 53(1), 11-22.

Arntz, Melanie, Michael Feil und Alexander Spermann (2003), Die Arbeitsangebotseffekte der neuen Mini- und Midijobs - eine ex-ante Evaluation, *Mitteilungen aus der Arbeitsmarkt- und Berufsforschung* 36, 271-290.

Aktuelle Diskussionspapiere/Arbeitspapiere

Arntz, Melanie, Simon Lo und Ralf Wilke (2007): Bounds Analysis of Competing Risks: a nonparametric evaluation of the effect of unemployment benefits on migration in Germany, *mimeo*, ZEW Mannheim.

Arntz, Melanie und Ralf Wilke (2006), *Unemployment Duration in Germany: Individual and Regional Determinants of Local Job Finding, Migration and Subsidized Employment*, ZEW Discussion Paper No. 06-092, Mannheim.

Arntz, Melanie, Stefan Boeters, Nicole Gürtzgen und Stefanie Schubert (2006), *Analysing Welfare Reform in a Microsimulation-AGE Model: The Value of Disaggregation*, ZEW Discussion Paper No. 06-076, Mannheim.

Arntz, Melanie (2006), *What Attracts Human Capital? Understanding the Skill Composition of Interregional Job Matches in Germany*, ZEW Discussion Paper No. 06-062, Mannheim.

Arntz, Melanie, Ralf Wilke und Henrik Winterhager (2006), *Regionenmatching im Rahmen der Evaluation der Experimentierklausel des § 6c SGB II: Methodische Vorgehensweise und Ergebnisse*, ZEW Discussion Paper No. 06-061, Mannheim.

Arntz, Melanie (2005), *The Geographical Mobility of Unemployed Workers*, ZEW Discussion Paper No. 05-34, Mannheim.

Beiträge in Sammelbänden

Arntz, Melanie, Peter Jacobebbinghaus und Alexander Spermann (2004), Minijobs, Midijobs und sozialversicherungspflichtige Beschäftigung in privaten Haushalten, in: Tobias Hagen, Alexander Spermann, Hartz-Gesetze - *Methodische Ansätze zu einer Evaluierung*, ZEW Wirtschaftsanalysen, Bd. 74, Baden-Baden (unter Mitarbeit von: Melanie Arntz, Andreas Ammermüller, Miriam Beblo, Martin Biewen, Bernhard Boockmann, Marco Caliendo, Bernd Fitzenberger, Nicole Gürtzgen, Reinhard Hujer, Peter Jacobebbinghaus, Friedhelm Pfeiffer, Andrea Weber, Ralf Wilke, Henrik Winterhager, Elke Wolf, Gaby Wunderlich), S. 171-189.

Arntz, Melanie und Alexander Spermann (2004), Nationale und internationale Konzepte in der Praxis, in: Thomas Klie, Alexander Spermann, *Persönliche Budgets - Aufbruch oder Irrweg?*, 1. Auflage, Ein Werkbuch zu Budgets in der Pflege und für Menschen mit Behinderungen, Hannover, S.16-39.

Arntz, Melanie und Alexander Spermann (2004), Probleme einer Einführung eines personengebundenen Pflegebudgets, in: Thomas Klie, Alexander Spermann, *Persönliche Budgets - Aufbruch oder Irrweg?*, 1. Auflage, Ein Werkbuch zu Budgets in der Pflege und für Menschen mit Behinderungen, Hannover, S. 222-250.

Mannheim, den 27.06.2007